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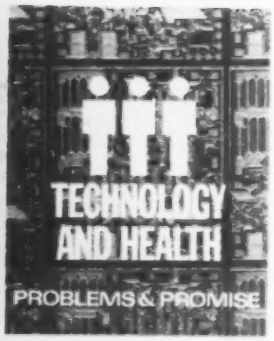

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A Cost-Effectiveness Analysis of Exercise as a Health Promotion Activity

EVRIDIKI I. HATZIANDREU, MD, MPH, JEFFREY P. KOPLAN, MD, MPH, MILTON C. WEINSTEIN, PHD, CARL J. CASPERSEN, PHD, AND KENNETH E. WARNER, PHD

Abstract: We used cost-effectiveness analysis to estimate the health and economic implications of exercise in preventing coronary heart disease (CHD). We assumed that nonexercisers have a relative risk of 2.0 for a CHD event. Two hypothetical cohorts (one with exercise and the other without exercise) of 1,000 35-year-old men were followed for 30 years to observe differences in the number of CHD events, life expectancy, and quality-adjusted life expectancy. We used jogging as an example to calculate cost, injury rates, adherence, and the value of time spent. Both direct and indirect costs associated with exercise, injury, and treating CHD were considered. We estimate that exercising regularly results in 78.1 fewer CHD events and 1,138.3 Quality Adjusted Life Years (QALYs) gained over the 30-year study period. Under our base case assumptions, which include indirect costs such as time spent in exercise, exercise does not produce economic savings. However, the cost per QALY gained of $11,313 is favorable when compared with other preventive or therapeutic interventions for CHD. The value of time spent is a crucial factor, influencing whether exercise is a cost-saving activity. In an alternative model, where all members of the cohort exercise for one year, and then only those who like it or are neutral continue, exercise produces net economic savings as well as reducing morbidity. (*Am J Public Health* 1988; 78:1417–1421.)

Introduction

Regular aerobic exercise is increasingly being viewed by health professionals as a key behavioral ingredient in reducing the risk of illness, particularly illness associated with heart disease. The "fitness craze" has emerged as an integral component of the broader concern with health promotion, with exercise explicitly recognized in the US Public Health Service's health promotion and disease prevention objectives for the nation.[1] In recent years the presumed cardiovascular benefits of exercise have received convincing empirical support.[2–4] Accordingly and increasingly, the medical profession has come to recognize and prescribe exercise as a preventive and rehabilitative therapy.

Many health promotion advocates emphasize also the economic benefits of exercise, in addition to its contribution to better health.[5] Because both the economic and medical dimensions of exercise are subjects of substantial interest, we examine the cost-effectiveness of exercise as a primary prevention medical technology. Our interests are in the health and economic implications of exercise as compared with its absence and in using exercise as an alternative to common medical interventions designed to address potential or existing heart disease. In addition, we examine the issue of whether exercise consumes more or less time than it saves in increased life expectancy.

Methods

We used decision analysis and cost-effectiveness analysis to compare a regular exercise regimen with its absence in cohorts of 1,000 35-year-old men, followed for 30 years. We defined exercise as leisure time physical activity that consumes 2,000 kcal per week. Jogging was used as an example to calculate costs, injury rates, adherence, and the value of the time spent. Exercise was considered to be a prescriptive medical intervention, analogous to chronic antihypertensive or hypoglycemic treatment. Cost-effective is a relative and subjective concept defined as "having an additional benefit worth the additional cost."[6] Beneficial programs that yield net cost savings are automatically cost-effective, but net cost saving is not necessary for a program to be judged cost-effective.

Using epidemiologic data, we estimated the number of fatal and nonfatal coronary heart disease (CHD) events that would be expected to occur over time in each cohort. We considered the consequences of CHD events to be their direct medical care cost and their effects on life expectancy and reduced quality of life. Thus, our measures of the health effectiveness of exercise were reductions in the numbers of CHD events, gains in life expectancy, and gains in quality-adjusted life expectancy. These effectiveness measures served as denominators in our cost-effectiveness ratios. The numerators of the ratio represent various measures of the difference between the incurred costs of the two strategies. We considered both direct costs and savings associated with an exercise regimen (direct costs, including equipment and clothing, as well as the medical cost of injuries, and savings consisting of the avoidance of treatment for CHD events) and indirect costs and benefits, respectively (time spent in the exercise program and healthy time gained as a result of CHD events prevented). Our outcome measures are cost (direct or direct plus indirect) per year of life saved, cost per quality-adjusted life year (QALY) saved, and cost per CHD death averted. All future costs and years of life saved were discounted at a real annual rate of 3 per cent to determine their present value. We chose to discount health benefits (future years of life saved and cases averted) to maintain a common perspective on the value of both future costs and future health effects. Costs were measured at 1985 price and wage levels.

Definitions and Rates

Health events for which rates were required in the analysis were exercise-related injuries, CHD events, and non-CHD mortality. Injury was defined as an adverse health effect severe enough for the injured person to seek medical consultation. We estimated the annual probability of such injury to be 5 per cent, with 17 per cent of those injured

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Editor's Note: See also related editorial p 1413 this issue.

quitting jogging permanently because of the injury.[7] We assumed that the rate of injuries and quitting remained constant throughout the 30-year study period.

CHD events were considered to be death (sudden or nonsudden), myocardial infarction, or other CHD (angina pectoris and coronary insufficiency). Age-specific incidence rates for men were derived from the Framingham Heart Study.[8,9] These rates, however, represent combined population rates, which include those who do and those who do not exercise. We calculated separate rates for exercisers and nonexercisers by assuming that 10 per cent of all men exercise as defined previously and that exercise reduces the risk of CHD by 50 per cent.[10] Age-specific death rates for men for 1980 from all other causes except CHD were used.

In both the exercise and nonexercise cohorts, we separated those who experienced nonfatal CHD into subgroups and tracked them to determine the number of additional deaths occurring during the follow-up period. Annual mortality rates in these subgroups of persons with CHD histories were based on data from the American Heart Association (age-specific CHD deaths and prevalence of CHD). Persons injured due to exercise and unable to continue jogging were considered "quitters" (dropouts). We assumed that they did not engage in other forms of aerobic activity and that they experienced the CHD rates of non-exercisers from the year of quitting.

Costs

Exercise—The direct cost of exercise, the cost of equipment (running shoes, clothes, etc.), is estimated at $100 annually. Considering exercise to be a prescriptive medical intervention, we have included the cost of counseling a patient to undertake an exercise regimen as a direct cost. We assumed that this counseling would occur during a routine periodic examination and would represent 10 per cent of a 45-minute visit. We estimated the average cost of the history and examination portion of a routine physician visit to be $75.

To determine the indirect cost, we assumed five hours per week, including exercise time (nine minutes per mile, 20 miles per week) and preparation time (24 minutes, five times a week). In deriving the monetary value of this time, we used the results of a telephone survey of six exercise experts who estimated the proportions of participants in an exercise program who like it (55 per cent), dislike it (35 per cent), or are neutral (10 per cent). We assumed that those who are neutral toward exercise would value their time at $4.50 per hour (half the 1985 hourly wage.[11]) Those who dislike exercise would value their time at $9 per hour, the average hourly wage, whereas those who like exercise would view it as costless, the "enjoyment value" exactly canceling out the opportunity cost of time expended. Multiplying these monetary values by the survey proportions produced an average value of $3.57 per hour of exercise for the cohort.

Injury—For the direct medical cost of injury, we assumed that an average office visit costs $40, and 50 per cent of the injured have an x-ray taken, at an average cost of $50. For the indirect cost, we estimated the average hours of work lost because of injury.[7] Work time lost is less than one hour per injured person per year, which then multiplied by the average hourly wage, yields an estimate of $8 per injured person per year.

Coronary Heart Disease—We used published direct and indirect lifetime cost estimates for CHD.[9] Direct costs are emergency assistance, hospitalization, follow-up care, including office visits, tests, and medications, and possible coronary angiography and coronary artery bypass grafting. Indirect costs are losses of earnings due to disability and premature death. All costs were adjusted from 1980 to 1985 dollars.

Measures of Effectiveness

To measure the health effects of exercise programs, we calculated the expected number of CHD events and the gain in years of life saved to allow comparison with other preventive or therapeutic measures.[12] The gains in years of life expectancy were adjusted for changes in the quality of life due to decreased morbidity from nonfatal CHD. We calculated QALYs by assuming that each year which follows the onset of a nonfatal form of CHD is equal to .8 of a healthy year.[12] We calculated QALYs for exercise-related injuries by assuming that the quality of the time spent while injured was .9 of an equivalent uninjured time period.

Sensitivity Analysis

We performed sensitivity analyses to test several of our assumptions. First, we varied the relative risk for a CHD event from 1.5 to 2.5. In addition, we varied the prevalence of those who exercise from 10 per cent to 20 per cent; the discount rate from 0 per cent to 5 per cent; the direct cost of exercise from $100 to $200; the indirect cost of exercise from $2 to $13.50; and the QALY weight from .7 to .9.

Alternative Model

We have also considered an alternative (voluntary) approach, which ultimately would target the exercise program at those who consider it to be enjoyable, or are at least neutral. We assumed that persons who start exercising do so for a one-year trial. After that, those who like it or are neutral continue, whereas those who do not like it quit permanently. We used probabilities of liking or disliking exercise derived from the expert survey and applied the same cost assumptions.

Analysis of Time Trade-Off

Recent critiques of analysis of the life expectancy benefits of exercise have commented that exercisers consume as much time in exercise as they gain in life expectancy.[13] To examine this issue and to provide another perspective, we determined the balance between time spent on exercise and injuries and that gained from CHD prevention. In effect, this approach constitutes a time cost-benefit analysis of exercise. We used the same assumptions as in the base case cost-effectiveness analysis.

Results

Health Benefits of Exercise

Under the base case assumptions there were 78.1 fewer CHD events in the exercise cohort than in the nonexercisers (Table 1). Furthermore, because of gains in life expectancy and quality of life, there were 1,138.3 QALYs saved in the same cohort (529.8 if discounted) (Table 2).

Costs

The expected costs under the two alternatives are displayed in Table 3. Accounting for the indicated direct and indirect costs of exercise, injury, and disease, exercise does not produce economic savings. It costs $740,000 more in direct costs and $6 million more in total (direct and indirect) costs for the cohort of exercisers.

TABLE 1—Expected Events Occurring between Ages 35–64 in Cohorts of 1,000 Exercisers and Non-Exercisers

Exercise	No Exercise
Total CHD 197	118.9
Myocardial infarcts 85.6	51.5
Other CHD 81.9	49.3
Deaths (initial) 29.5	18.1
Additional Deaths* 26.7	14.2
Total CHD Deaths 56.2	32.3

*Deaths from CHD in cohort members whose initial event is MI or other CHD.

Cost-Effectiveness Ratics

Estimates of the cost-effectiveness of exercise in preventing CHD are presented for four different effectiveness measures: QALYs gained, years of life gained, CHD cases averted, and CHD deaths averted (Table 4). The costs per QALY gained when direct and total costs are considered are $1,395 and $11,313, respectively.

Sensitivity Analysis

Using discount rates of 0 per cent and 5 per cent produces total costs/QALY of $8,376 and $12,623, respectively (Figure 1). Using a relative risk for CHD for nonexercisers of 2.5 produces a cost/QALY of $6,596, whereas a relative risk of 1.5 produces a cost/QALY of $18,685. A prevalence rate of exercise of 20 per cent in the general population results in little change in the cost/QALY, $9,654. When QALY weights of .7 and .9 are used, the costs/QALY are $12,090 and $10,352, respectively. Doubling the direct costs of exercise to $200 increases the cost/QALY to $14,342. When the indirect cost of exercise is varied from the base case estimate of $3.57 per hour, the cost/QALY varies from no cost at $2 per hour ("break-even" point) to more than $24,061 per QALY when an average hourly wage of $9 is used (Figure 2).

Voluntary Approach to Exercise

With this alternative, voluntary approach, a gain of 345.4 QALY is expected for those who continue exercising (i.e., those who like it and those who are neutral). The results are presented in Tables 5 and 6. The health benefits are less than the QALYs gained in the main analysis (350.1 vs 529.8 QALYs); however, this voluntary approach to exercise yields net economic savings even when indirect costs are included ($23.13 million for exercisers vs $25.79 million for nonexercisers). People who do not like exercise quit and thus do not incur the cost of $9 per hour.

We determined the incremental cost and benefit of coercing those who simply do not like exercise to engage in jogging. The marginal cost-effectiveness ratio for the coerced population is $48,775 per QALY.

Analysis of Time Trade-off

When time is discounted, like any other cost or benefit, the cohort of the 1,000 men aged 35 who exercise regularly for the next 30 years spends a total of 522.7 years in exercising, whereas 530 years are gained by preventing CHD. The "benefit-cost" difference is positive by 7.3 years, representing a net gain of 2.7 days per member of the cohort. Without discounting, there are 730 years spent and 1,030 years gained, a "benefit-cost" difference of 300 years, or almost 110 days per cohort member.

TABLE 2—Years of Life Lost in 1,000 Exercisers and 1,000 Nonexercisers, Ages 35–64

	Not Discounted		Discounted*	
	Exercise (N=1000)	No Exercise (N=1000)	Exercise (N=1000)	No Exercise (N=1000)
Due to CHD				
from Death	672.2	1190.5	276.9	492.1
from morbidity (QALY)	899.9	1539.9	449.8	778.2
Due to Jogging Injury	20.0	—	13.8	—
Total	1592.1	2730.4	740.5	1270.3

*Present values at 3 per cent per annum.

TABLE 3—Expected Costs ($ Million) among 1,000 Exercisers and Non-Exercisers*

	Exercise (N=1000)			No Exercise (N=1000)		
	Direct	Indirect	Total	Direct	Indirect	Total
Equipment	1.670	—	1.670	—	—	—
Exercise time	—	15.450	15.450	—	—	—
Injury	.052	.006	.058	—	—	—
Disease	1.410	13.200	14.610	2.39	23.4	25.79
Total	3.132	28.656	31.788	2.39	23.4	25.79

*Present values at 3 per cent per annum.

Discussion

Previous studies have concluded that other disease prevention technologies are cost-effective relative to therapeutic health care expenditures,[14] and sometimes even produce net economic savings, e.g., fluoridation[15] and certain immunizations.[16,17] Recently, however, concern has been expressed that, in some instances, preventive measures for chronic illnesses may be less cost-effective than those for acute conditions, and further that preventive measures are not invariably more cost-effective than curative therapy.[18]

On the basis of our estimates and assumptions, exercise *is* a cost-effective approach to lowering the risk of CHD. Exercise costs $11,313 per QALY saved, including the cost of time spent, and $1,395 in direct costs, excluding indirect costs and benefits. These figures compare favorably with other published cost-effectiveness studies of CHD interventions.[12] Coronary artery bypass graft surgery with left main disease costs $5,000 per QALY when two-vessel disease is present. With mild angina the cost per QALY becomes $40,000. Treating hypertension varies from $25,000 per QALY (diastolic pressure of 95–104 mm Hg), to $65,000 per QALY (diastolic pressure 90–94 mm Hg). It should be

TABLE 4—Cost-Effectiveness Ratios*

	Direct Cost ($)	Total Cost ($) (Direct & Indirect)
Cost/QALY Gained	1,395	11,313
Cost/Year of Life Gained	3,433	27,851
Cost/CHD Case Averted	9,462	76,760
Cost/CHD Death Averted	30,920	250,836

*All costs and effectiveness measures discounted at 3 per cent per annum.

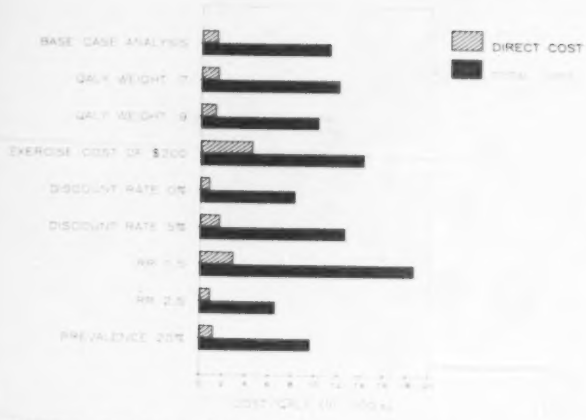

FIGURE 1—Sensitivity Analysis

emphasized, however, that these studies include direct costs only. It is conceivable that if these other CHD interventions saved a sufficient amount of productivity, then the inclusion of indirect benefits and costs would make these interventions more cost-effective.

The other cost-effectiveness studies of CHD interventions use only direct costs and savings, as is customary in cost-effectiveness analysis. We consider the indirect costs of time spent in exercise to be integral to this health intervention. An alternative approach to expressing cost-effectiveness could apply both direct and indirect costs but only direct benefits (our base case uses both direct and indirect benefits). This method results in a cost of $30,575 per QALY saved, or $75,232 per year of life saved. For the voluntary approach, exercise, which was cost saving when both direct and indirect costs and benefits were used, no longer saves money but is cost-effective at $9,180 per QALY saved, when total costs and only direct benefits are used. Thus even using this conservative approach, exercise remains comparable in cost-effectiveness to other CHD interventions. Significantly, evaluations of the costs of treatment of morbid conditions typically do not quantify the nonpecuniary costs of the disease and its treatment (i.e., the costs of pain and suffering).

The indirect cost of exercise is the most important

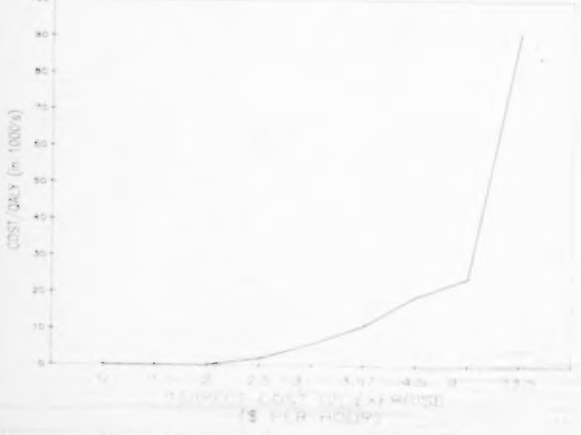

FIGURE 2—Relation of Indirect Cost of Exercise to Cost/QALY

TABLE 5—QALYs Lost under the Noncoercive Exercise Program*

	Exercise	No Exercise
Due to Death	350.1	492.1
Due to Disease	561.2	778.2
Due to Injury	9.0	0
Total	920.3	1270.3

*Present values at 3 per cent per annum.

variable in influencing the results of this analysis. If the indirect cost (time spent in exercise) were valued as "no cost", then exercise would be cost-saving. However, we treat exercise time as having opportunity cost. Although the wage rate is commonly used in economic studies as the value of leisure time, we used one-half the wage rate as our basic measure of this opportunity cost. Both theory and empirical analysis suggest that the opportunity cost of the marginal leisure hour is less than the wage rate. We have used the estimate of half the wage rate produced by one of the few studies to examine this issue empirically.[19] Note that our treatment of the indirect cost of time for those who enjoy exercise is undoubtedly conservative for many people who are (or would be) willing to pay to exercise (as demonstrated by memberships in health clubs, for example). To these people, the benefits of exercise—immediate and perceived for the future—give exercise a net positive economic value.

Our results support an approach in which physicians would strongly urge patients to begin an exercise program on a trial basis for one year, with the expectation that only those persons who find it pleasurable or those who are neutral would be encouraged to continue. Such an approach was found to yield net cost savings, whereas an all-or-nothing approach in which "exercise-haters" are forced (in the model) to comply was cost-ineffective for this group ($48,775/QALY) compared with other methods of coronary heart disease prevention ($5,000 to $30,000 per QALY).

Our study focused on only cardiovascular health benefits of exercise. Other demonstrated health benefits of exercise include weight control, stress reduction, smoking cessation, and assistance in diabetes and hypertension management.[20] Quantification and inclusion of these benefits would improve the relative cost-effectiveness of exercise, possibly considerably.

Other forms of exercise also provide aerobic cardiovascular benefits. If used in a similar analysis, walking, swimming, and bicycling would require changing several variables. For example, walking might be a more enjoyable pastime for certain participants, thus lowering the indirect costs of exercise, but walking would require more time to achieve the same cardiovascular benefit, thereby increasing

TABLE 6—Costs under the Noncoercive Exercise Program ($ Million)*

	Exercise			No Exercise		
	Direct	Indirect	Total	Direct	Indirect	Total
Equipment	1.117	—	1.117	—	—	—
Exercise time		2.80	2.800	—	—	—
Injury	.004	—	.004	—	—	—
Disease	1.780	17.43	19.210	2.39	23.4	25.79
Total	2.901	20.23	23.131	2.39	23.4	25.79

*Present values at 3 per cent per annum.

the indirect costs. It would likely also reduce the direct costs of injury caused by the exercise.

Our study focused exclusively on 35-year-old men followed for 30 years. Further studies in exercise cost-effectiveness could consider women, various racial, ethnic or socioeconomic groups, other forms of exercise, and other health benefits. Adherence to an exercise regimen would affect the results of this analysis just as would adherence to a drug therapy. Partial adherence would reduce the cost (indirect) of exercise and the injury rate but also the cardiovascular benefits conferred by exercise. Thus, additional aspects that could be studied would be to vary adherence and the number of years of participation in an exercise program.

In another analysis of the health benefits of exercise, Paffenbarger *et al*,[21] estimated that exercise adds 2.15 years to the life expectancy. If we change some of the assumptions in our analysis (no discounting, full adherence, and no QALY adjustment), the total years gained for the cohort are 1,500, which is 1.5 years per person. The difference can be attributed to the fact that Paffenbarger, *et al*, looked at mortality from all causes from a very specific cohort (Harvard alumni), whereas we looked at increases in life expectancy by preventing CHD only, using national estimates.

To perform this analysis we have had to make several assumptions and use the best available data. The analysis can be improved in the future as more knowledge is gained on exercise and CHD, and then applied in a systematic and analytic framework.

For persons who enjoy the immediate pleasure and benefits of exercise, this activity is probably also a cost-effective measure in reducing risks of CHD. For others who find exercise a less pleasant activity, its cost-effectiveness is comparable to that of some drug therapies and surgical interventions.[12] Most persons logically would prefer to prevent a morbid condition rather than develop it and have it treated. However, they may require more immediate satisfaction to undertake and maintain a regular exercise regimen than the knowledge that it will reduce their long-term risk of CHD or that it is "cost-effective."

ACKNOWLEDGMENTS

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Divergence of the Recent Trends in Coronary Mortality for the Four Major Race-Sex Groups in the United States

CHRISTOPHER SEMPOS, PHD, RICHARD COOPER, MD, MARY GRACE KOVAR, DRPH, AND MARILYN MCMILLEN, PHD

Abstract: Since 1976 there has been a leveling off or slowdown in the rate of decline in coronary heart disease (CHD) mortality. The age-adjusted absolute annual rate of decline in CHD mortality rates during 1968–75 (Δ rate/100,000 population/year) was virtually identical for White males (−7.54), Black males (−7.85), and Black females (−7.20), and somewhat lower for White females (−4.25). During 1976–85, however, the secular trends diverged considerably. Age-adjusted rates continued to decline at the same annual rate for White males, while the decline was approximately half as steep for the other thre race-sex groups. During 1976–85 there was also a leveling off in the average annual per cent change in age-adjusted CHD mortality for Black males and females and White females when compared to 1968–75, while there was no change for White males. As a result, more than 40,000 White and Black females and Black males died of CHD in 1985 than would have died if CHD rates would have continued to decline at the 1968–75 trends. All comparisons were based on a reclassification of cause-of-death codes to maximize comparability between the 8th and 9th Revisions of the International Classification of Disease. These results suggest that the factors which have led to the continued decline in coronary heart disease may not have influenced all the demographic groups in this country equally over the last decade. (*Am J Public Health* 1988; 78:1422–1427.)

Introduction

Rapidly changing patterns of mortality from coronary heart disease (CHD) have been recognized in a wide range of countries over the last three decades.[1–3] Beginning in the mid-1960s, a tendency has been observed for countries with the highest rates to experience a rapid and consistent decline in mortality, most notably in the United States, Canada, Australia, New Zealand, and Finland, while Eastern Europe and the Soviet Union, which entered this period with lower rates, have seen a marked upturn in death rates from CHD.[3–6]

Although preventive campaigns and changes in the prevalence of risk factors for CHD may have contributed to the decline in the US, it has not been possible to provide precise estimates of all the factors affecting the downward trend in CHD mortality rates.[7–9] The prevalence and changes in the prevalence of the coronary risk factors have not been the same among the four major race-sex groups of the US[8,10–12]; it is therefore important to examine the secular trends in CHD mortality by sex and race to determine whether the changes are the same for the four groups.

Around 1980 it appeared that the decline in mortality for adults from all causes combined had slowed in the US, with a brief increase recorded for Blacks.[13] Since CHD is by far the most important single disease category determining adult mortality rates in the US, and it has been the primary component of significant secular trends in all-causes mortality, we examined trends in CHD mortality for the four major race-sex groups in the US with the hope of gaining insight into the two important questions: Has the decline in CHD mortality continued at the same rate in the 1980s? Has it continued at the same rate for each of the four largest race-sex groups?

Methods

Data Sources and Definitions

Age-adjusted and age-specific CHD mortality data for the four major race-sex groups were obtained from vital statistics data tabulated by the National Center for Health Statistics for the years 1968–85. Age-adjusted rates were computed by the direct method using the total US population of 1940 as the standard. During 1968–78 deaths were coded using the 8th revision of the International Classification of Diseases Adapted for Use in the United States (ICDA 8) and during 1979–85 they were coded using the 9th Revision of the ICD (ICD 9). We adjusted for the lack of comparability in coding rules between the two revisions by including in CHD for 1979–85 those ICD 9 categories that, for the most part, would have been coded in CHD categories in ICDA 8.[14] As a result, CHD during 1968–78 was defined using the category numbers for Ischemic Heart Disease (ICDA Nos. 410–413) and during 1979–1985 it was defined as the sum of Ischemic Heart Disease (ICD Nos. 410–414), Hypertensive Heart Disease (ICD No. 402), and Cardiovascular Disease Unspecified (ICD No. 429.2) (See Appendices Tables 1 and 2).

Because there was no break in comparability for Acute Myocardial Infarction (MI) and Diseases of Heart between the two revisions those two categories were used to confirm the trends for CHD.[14] Acute MI, a category within CHD, was coded using category No. 410 during both revisions. Diseases of Heart, a more general category which includes CHD, was defined during the 8th Revision as ICDA Nos. 390–398, 402, 404, 410–429 and during the 9th Revision as ICD Nos. 390–398, 402, 404–429.

Statistical Methods

Linear regression was used to compare the slopes for both the *absolute* and *relative* rates of change in CHD mortality.[15–17] The *absolute* rate of change in CHD mortality was calculated as the change in the age-adjusted mortality rate/100,000 population/year. A log-linear model, which assumes that there is a constant proportional or *relative* change, was used to calculate the average annual per cent change.[17] Data for Figure 1 and Tables 1, 2 and 4 were calculated using the entire age range, and for Table 3 they were calculated only for people ages 35–74.

Because of the prior work that found evidence for a slowdown in the decrease in all cause mortality,[13] we anticipated that the slope or rate of decline in CHD during 1976–85 would be less than during 1968–75, i.e., that mortality rates for CHD decreased at a slower rate during 1976–

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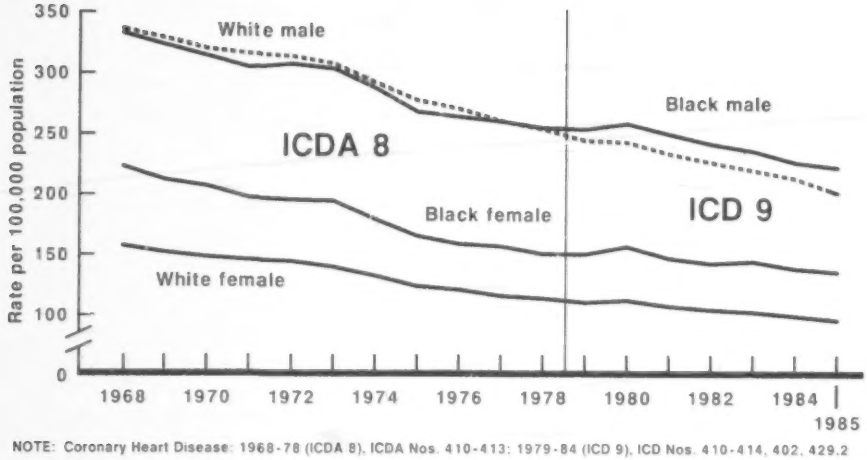
Editor's Note: See also related editorial p 1415 this issue.



NOTE: Coronary Heart Disease: 1968-78 (ICDA 8), ICDA Nos. 410-413; 1979-84 (ICD 9), ICD Nos. 410-414, 402, 429.2
SOURCE: Prepared by authors using data from NCHS, Division of Vital Statistics

FIGURE 1—Trends in age-adjusted Coronary Heart Disease for all ages, United States, 1968–84.

85 than during 1968–75. But, since it is important to monitor any change in mortality trends a two-sided t-test was used in comparing the slopes from the two periods. Statistical significance was set at the 0.05 level with a critical t-value (df = 14) of 2.145. To examine the effects of our reclassification of CHD during ICD 9, the same procedures were used to compare slopes for the periods 1968–78 and 1979–85. However, because we were examining trends in CHD mortality over a relatively short period of time, we emphasize consistency in the data more than statistical significance.

To investigate the impact of our recategorization of the mortality data coded under the 9th Revision on the analysis of CHD trends, comparisons were also made using first the original 9th ICD classification for Ischemic Heart Disease (ICD Nos. 410–414) and then adding incrementally Hypertensive Heart Disease (ICD No. 402) and Cardiovascular Disease Unspecified (ICD No. 429.2) (Table 1). The values for either absolute or relative rate of change in CHD mortality were not changed by including the other two categories. Addition of Hypertensive Heart Disease and Cardiovascular Disease Unspecified to Ischemic Heart Disease minimized the discontinuity between the two ICD periods without affecting the trends during 1979–85.

Projected age-adjusted CHD mortality rates for 1985 were determined using the race-sex specific regression equations for 1968–75. Excess CHD mortality associated with the slowing up in the rate of decline after 1975 was estimated by comparing the observed and the projected age-adjusted rates for 1985. The observed age-adjusted rate minus the projected one divided by the observed was used to define the percentage excess in age-adjusted mortality. The excess number of CHD deaths in 1985 was calculated by multiplying the percentage excess in the age-adjusted rates by the observed number of CHD deaths for the year 1985.

TABLE 1—Absolute and Relative Changes in Coronary Heart Disease Mortality During the Period of ICDA 8-1968–78 and the period of ICD 9-1979–85 by ICD Category Numbers, Race and Sex: United States, 1968–85

Periods	ICD Code Numbers	White		Black	
		Males	Females	Males	Females
		Absolute Rate of Change per 100,000 Population per Year			
1968–78	410–413	−8.45	−4.51	−8.02	−7.29
1979–85	410–414	−6.87	−2.37	−5.17	−2.18
	410–414,402	−6.95	−2.51	−5.49	−2.60
	410–414,402,429.2	−7.15	−2.65	−6.16	−2.96
		Average Annual Per Cent Change			
1968–78	410–413	−2.83	−3.30	−2.72	−3.89
1979–85	410–414	−3.37	−2.59	−2.84	−2.00
	410–414,402	−3.33	−2.63	−2.69	−2.07
	410–414,402,429.2	−3.12	−2.50	−2.53	−2.02

NOTES: All rates were age-adjusted by the direct method using the total US population of 1940 as the standard.
SOURCE: Prepared by authors using data from the NCHS, Division of Vital Statistics.

Results

CHD mortality rates (age-adjusted) have declined in the US for all four of the major race-sex groups throughout 1968–85 (Figure 1). A break or leveling off in the trend lines for White females, and for Black males and females appears around 1976, while there is no discernible break or leveling off for White males. (The rises in mortality discernible in 1972–73 and 1980 are related to increases in CHD mortality during influenza epidemics.)

With Black males and females and White females, both the absolute and the relative rates of decline after 1975 were slower than during the period 1968–75 (Table 2). Black females experienced the greatest leveling off in CHD mortality in both absolute and relative terms, followed by White females, and then by Black males. White males did not experience a slowdown; in fact there was evidence of an increase in their rate of decline.

We replicated the analysis shown in Table 2 for ages 35 through 74 years (Table 3). The age-adjusted data for this restricted age range support the hypothesis of a slowdown in the rate of decline in CHD mortality after 1975 for White and Black females and Black males, although, the results for Black males, were not "statistically significant". Furthermore, results consistent with Table 2 were found when

TABLE 2—Change in Age-adjusted Mortality Rates for Coronary Heart Disease for all Ages between 1968–75 and 1976–85: United States, 1968–85

Periods	White		Black	
	Males	Females	Males	Females
	Absolute Rate of Change per 100,000 Population per Year			
1968–75	−7.54	−4.25	−7.85	−7.20
1976–85	−7.16	−2.52	−4.57	−2.41
t	(0.53)	(4.40)	(3.05)	(6.47)
	Average Annual Per Cent Change			
1968–75	−2.42	−2.97	−2.57	−3.64
1976–85	−2.99	−2.30	−1.86	−1.62
t	(−2.24)	(2.07)	(1.76)	(4.56)

NOTES: All rates were age-adjusted by the direct method using the total US population of 1940 as the standard; Coronary Heart Disease: 1968–78 (ICDA 8) ICDA Nos. 410–413; 1979–85 (ICD 9) ICD Nos. 410–414, 402, 429.2; t—statistic for two-sided t-test on the difference between the two specified periods, $t_{0.05}$ (d.f. = 14) = 2.145.

SOURCE: Prepared by authors using data from the NCHS, Division of Vital Statistics.

age-adjusted trends for Acute MI and Disease of Heart were compared between 1968–75 and the 1976–85 (data available from authors).

Age-specific trends are also shown in Table 3. Black women experienced an absolute and relative slowdown across all age groups, while the slowdown among White women appeared to be present in only the older age groups. An absolute slowdown was present in all age groups for Black men, while the relative changes in CHD mortality do not follow a clear pattern.

The excess mortality associated with the slow down in CHD mortality rates is shown in Table 4. Black females had a mortality rate from CHD in 1985 that was 27 per cent higher than the rate projected from the 1968–75 trend; White females and Black males had rates that were approximately 10 per cent higher. Based on the percentage changes in age-adjusted mortality, there were an estimated 40,000 more CHD deaths for White and Black females and Black males in 1985 than there would have been had the 1968–75 trend continued. The majority of these deaths (30,000) were of White females.

Discussion

One of the more remarkable features of the decline in CHD mortality that began in the US in the mid-1960s was its virtually simultaneous onset and similar relative magnitude of decline in the four major race-sex groups, and across the entire age range where deaths from this disease occur.[1,2] The analysis presented here suggests that this pattern of decline began to change significantly in the mid-1970s and, although age-adjusted CHD mortality rates continued in a downward trend for each of the four demographic groups, the rapid rate of decline observed from 1968–75 has persisted only for White males.

Regardless of whether we classified deaths during 1979–85 to minimize the effect of changing from ICDA 8 to ICD 9, used 1979–85 or 1976–85 as the second time period, or used absolute or per cent changes, the results were the same. The rate of decline has not slowed for White males and has slowed for females of both races and to a lesser extent for Black males.

We do not believe that the divergence of CHD trends is a result of coding changes that were instituted in 1979 with the introduction of ICD 9. The divergence began around 1975–76, three years before the introduction of ICD9, and the trends for 1976–85 were very similar to those for 1979–85. Moreover, the results for all ages (Table 2) were similar to the results where the analysis was restricted to persons ages 35–74 years (Table 3). Finally, the same results were found for Acute MI and Diseases of Heart, two disease categories with no discontinuity between the two ICD periods.

The major difficulty in evaluating CHD trends in the US, in the absence of long-term community-based myocardial infarction registries, remains the validity of the death certificate.[19–23] We can find no evidence that secular trends over the last two decades have been influenced by significant changes in patterns of reporting cause of death by medical practitioners. More important is whether a systematic bias in reporting for the different race-sex groups[24] exists and has changed, or whether changes in the classification system alter the final code assigned. There is evidence, albeit less than satisfactory, that Blacks have a higher proportion of out-of-hospital deaths from CHD[25–28a] than Whites. Comparison of death certificate data and hospital admissions in at least three communities support this contention,[25–27] and findings from the pilot phase of the Community and Cohort Surveillance Project also suggest that a higher proportion of CHD deaths among Blacks occur out of the hospital.[29] It has also been demonstrated recently that the delay before receiving medical care during the acute phase of MI is longer among a group of Blacks seen at a large urban municipal hospital than previously reported for Whites.[30] None of these studies, addressed the issue of change over time in coding these deaths, however.

Why are the death rates from CHD no longer falling at the same rate for all the four race-sex groups? Any attempt to analyze the decline in CHD mortality must take into account changes in the incidence of the disease and case-fatality rates.[7] A variety of community-based studies have suggested that the major factor in the US CHD trends during 1968–80 has been the decrease in incidence.[20,22,31–33] However, recent reports have suggested that the incidence of MI, as estimated by hospital admissions, may have increased in recent years;[34,35] although this could be an artifact resulting from more people with acute MIs reaching the hospital alive. Conflicting evidence also exists about the improvement in acute case-fatality rates from MI and the long-term prognosis. Data from two community surveys in Minnesota[32,36,37] from the National Hospital Discharge Survey,[35] and from DuPont[31] suggest a fall in case-fatality rates, while summary data for hospitals in Boston[38] and the Health Insurance Plan of New York[39] suggest no change.

If the leveling off in CHD mortality which we have observed is a result of changes in the trends in incidence of CHD then this should be reflected in changes in the distributions of major CHD risk factors, such as serum cholesterol, overweight, hypertension, and cigarette smoking. The best sources of data on risk factor levels in the US population—the National Health and Nutrition Examination Survey (NHANES) and the National Health Interview Survey—are inconsistent.[8,10,40,41] Between the midpoints of NHANES I (1974) and NHANES II (1978), mean serum cholesterol levels were estimated to have decreased by approximately 3–4 mg/100 ml for White and Black men and Black women, and 1 mg/100 ml for White women.[10] During that same time period, the prevalence of overweight remained unchanged for Black and

TABLE 3—Change in Mortality Rates for Coronary Heart Disease† for Persons Ages 35–74 Years during 1968–75 and 1976–84: United States, 1968–85

	White Men		White Women		Black Men		Black Women	
Age Group	1968–75	1976–85	1968–75	1976–85	1968–75	1976–85	1968–75	1976–85
	Absolute Rate of Change per 100,000 Population per Year							
35–74*	−14.52	−14.11	−6.26	−3.39#	−14.96	−11.27	−13.36	−5.93#
35–44	−3.01	−2.40#	−0.48	−0.41	−6.43	−2.96#	−3.94	−1.45#
45–54	−8.21	−9.61	−1.69	−1.50	−9.27	−9.85	−8.68	−5.37#
55–64	−22.39	−22.11	−7.10	−4.50#	−26.17	−17.52	−24.06	−6.96#
65–74	−50.02	−45.42	−32.60	−14.73#	−34.68	−28.28	−34.07	−18.44#
	Average Annual Per Cent Change							
35–74*	−2.62	−3.41	−3.26	−2.36#	−2.48	−2.36	−3.65	−2.27#
35–44	−3.82	−4.57	−3.21	−4.13	−5.07	−3.33#	−6.45	−4.24#
45–54	−2.54	−4.08	−2.51	−2.90	−2.25	−3.00	−4.19	−3.72
55–64	−2.57	−3.42	−2.75	−2.25	−2.77	−2.28	−4.22	−1.70#
65–74	−2.55	−3.04	−3.64	−2.24#	−1.86	−1.91	−2.64	−1.93

*Rates were adjusted by the direct method using the total US population of 1940 as the standard.
†Coronary Heart Disease: 1968–78 (ICDA 8) ICDA Nos. 410–413; 1979–85 (ICD 9) ICD Nos. 410–414, 402, 429.2.
#$p < 0.05$, for two-sided t-test $t_{0.05}$ (d.f. = 14) = 2.145 between 1968–75 and 1976–85.
SOURCE: Prepared by authors using data from the NCHS, Division of Vital Statistics.

White women and increased somewhat for Black and White men.[41] The prevalence of definite hypertension (BP ≥160/95 mm Hg and/or currently taking antihypertensive medication) during the period from 1974 to 1978 increased slightly (1–2 per cent) for White males and females and decreased by approximately 3–4 per cent for Black males and females.[42] Finally, between 1965 and 1976 rates of current smoking have decreased by 10 per cent for both Black and White men while there have only been small changes for Black and White women.[41]

Considering simultaneous changes in serum cholesterol, smoking, and blood pressure between 1974 and 1978 yields a somewhat different picture, however.[40] In this case, the impact of the changes in these three risk factors for persons ages 35–74 years was assessed by comparing the percentage decrease in observed CHD mortality between midpoints of NHANES I (1974) and NHANES II (1978) with the expected percentage change in CHD. The expected change in CHD mortality was estimated based on each individuals' levels of the three risk factors in both surveys by using a multiple logistic model derived from the Framingham data. The changes in risk factors explained 50 per cent or more of the observed decrease in CHD mortality between the two surveys. Thus, on a national level, small simultaneous changes in multiple risk factors can have important effects.

Other factors, such as competing cause effects and access to medical care, could alter these relationships. Additionally, such factors as economic trends and socioeconomic status or social class may also play a substantial role in changes in CHD mortality trends.[43–52]

Our purpose in this study was to document the divergence in CHD mortality trends and the slowing of the rate of decline for three of the four major race-sex groups in the United States. The public health importance of the slowing in rates of decline is evidenced by the approximately 40,000 excess CHD deaths of White females and Black males and females that would not have occurred if mortality rates had continued at the rates observed during 1968–75. The results emphasize the need for increased efforts aimed at primary and secondary prevention and access to appropriate treatment in Blacks and in White females, while maintaining and improving upon the gains already made in White males.

TABLE 4—Comparison of Observed CHD Mortality for all Ages in 1985 with Mortality Projected from Trends during 1968–75

		Age-adjusted Rate (per 100,000 Population)		Difference in Mortality	
Race	Sex	Observed	Projected	Rate (%)	Number of deaths
White	Males	204.4	210.1	−2.8	−8,327
	Females	96.6	85.9	+11.1	+29,973
Black	Males	222.7	199.5	+10.4	+3,000
	Females	136.4	99.8	+26.8	+7,496

NOTES: All rates were age-adjusted by the direct method using the total US population of 1940 as the standard; Coronary Heart Disease: 1968–78 (ICDA 8) ICDA Nos. 410–413; 1979–85 (ICD 9) ICD Nos. 410–414, 402, 429.2; Difference in mortality: Rate (%) = (observed − projected/observed) × 100; Number of deaths = (observed − projected).
SOURCE: Prepared by authors using data from the NCHS, Division of Vital Statistics.

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(continued)

APPENDIX TABLE 1--Age-adjusted Mortality Rates per 100,000 Population for Ischemic Heart Disease: United States 1968–85

	White		Black	
Years	Males	Females	Males	Females
ICDA 8 Category Nos. 410–413				
1968	336.6	157.6	332.9	223.2
1969	329.1	152.4	323.6	212.3
1970	320.3	148.5	314.5	207.1
1971	316.4	146.2	305.1	197.6
1972	314.2	144.8	307.9	195.6
1973	308.1	140.1	303.9	194.7
1974	292.4	132.9	287.8	179.6
1975	277.9	124.3	268.4	165.8
1976	271.1	121.4	264.5	159.2
1977	261.2	116.5	260.8	157.5
1978	254.2	114.5	255.4	151.3
ICD 9 Category Nos. 410–414				
1979	219.8	96.0	190.5	111.0
1980	218.0	97.4	196.0	116.1
1981	209.1	93.2	188.8	110.1
1982	203.0	90.9	182.2	106.9
1983	195.7	89.0	175.8	106.8
1984	187.0	86.0	168.5	102.6
1985	180.8	82.9	164.9	100.8

NOTE: All rates were age-adjusted by the direct method using the total US population of 1940 as the standard.
SOURCE: NCHS, Division of Vital Statistics.

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APPENDIX TABLE 2—Age-adjusted Mortality Rates per 100,000 Population for Hypertensive Heart Disease and Cardiovascular Disease Unspecified, and their Sum with Ischemic Heart Disease by Race and Sex: United States, 1979–85

	White		Black		White		Black	
Year	Males	Females	Males	Females	Males	Females	Males	Females
	Hypertensive Heart Disease ICD No. 402				Cardiovascular Disease Unspecified ICD No. 429.2			
1979	5.0	4.2	22.6	18.1	20.1	11.0	41.1	21.9
1980	5.0	4.1	21.8	17.5	20.6	11.5	41.2	23.8
1981	5.0	3.9	21.9	16.5	20.0	11.0	39.6	21.1
1982	4.8	3.8	21.6	16.1	19.2	10.5	38.3	20.4
1983	4.8	3.6	21.6	16.0	20.0	10.8	39.0	22.3
1984	4.6	3.5	20.7	15.7	19.5	10.6	37.7	20.8
1985	4.6	3.4	20.5	15.5	19.0	10.3	37.3	20.1
	Ischemic Heart Disease + Hypertensive Heart Disease				Ischemic Heart Disease + Hypertensive Heart Disease + Cardiovascular Disease Unspecified			
1979	224.8	100.2	213.1	244.9	111.2	254.2	151.0	
1980	223.0	101.5	217.8	133.6	243.6	113.0	259.0	157.4
1981	214.1	97.1	210.7	126.6	234.1	108.1	250.3	147.7
1982	207.8	94.7	203.8	123.0	227.0	105.2	242.1	143.4
1983	200.5	92.6	197.4	122.8	220.5	103.4	236.4	145.1
1984	191.6	89.5	189.2	118.3	211.1	100.1	226.9	139.1
1985	185.4	86.3	185.4	116.3	204.4	96.6	222.7	136.4

NOTES: All rates were age-adjusted by the direct method using the total US population in 1940 as the standard; Ischemic Heart Disease: 1979–84 (ICD 9) ICD Nos. 410–414.
SOURCE: Prepared by authors using data from NCHS, Division of Vital Statistics.

Rural Hospital Closures May Harm the Elderly

The closure of rural hospitals may hold serious implications for the very old living in rural areas of the United States. According to a research study published in the July issue of the *Journal of Rural Health*, such closures could have a critical impact on the rural elderly. The study, sponsored by the National Center for Health Services Research (NCHSR), analyzed distances traveled for hospital care by rural area residents in New York State. Few residents in the 75+ age group traveled beyond the county line for care. Most were hospitalized locally. Only 18 per cent crossed a county line to get care, compared with 31 per cent of those younger than age 75. Overall, 29 per cent of the State's rural inhabitants went outside their home county for hospital care. Two-thirds of this group went to urban medical centers.

The study, entitled "Patterns of Travel for Rural Individuals Hospitalized in New York State: Relationships between Distance, Destination, and Casemix," was conducted by Christopher Hogan, PhD, a former researcher at NCHSR. Reprints of the study are available from NCHSR Publications and Information Branch, Parklawn Building, Room 18-12, Rockville, MD 20857. Tel: (301) 443-4100.

Physical Activity and Coronary Heart Disease among Asymptomatic Hypercholesterolemic Men

(The Lipid Research Clinics Coronary Primary Prevention Trial)

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Abstract: We examined the relation between reported regular strenuous exercise or hard physical labor and the incidence of coronary heart disease (CHD) death and nonfatal myocardial infarction among 1,533 hypercholesterolemic men aged 35–59 years who were in the placebo group of the Lipid Research Clinics Coronary Primary Prevention Trial. The mean follow-up of the cohort was 7.4 years. The men were free of clinical heart disease at entry; men with an abnormal resting electrocardiogram or graded exercise test also were excluded.

Regular physical activity was not associated with the incidence of CHD (RR = .94, 95% CI = .68, 1.38) in this study population. Adjustment by the proportional hazards model for age, low-density lipoprotein cholesterol, smoking, family history of CHD, and occupation did not alter this finding. This observation suggests that reported regular physical activity may not be related to the risk of coronary heart disease among asymptomatic, hypercholesterolemic, middle-aged men. (*Am J Public Health* 1988; 78:1428–1431.)

Introduction

There is mounting evidence that regular strenuous physical activity is associated with a reduced risk of coronary heart disease,[1–4] but whether this relation applies to hypercholesterolemic men remains unresolved. Hypercholesterolemic men might particularly benefit from regular strenuous activity, if regular activity significantly altered lipoprotein levels[5] or mitigated their adverse effects. On the other hand, hypercholesterolemic men might benefit little or not at all from regular physical activity, if the cardiac benefits do not outweigh the risks of strenuous activity[6] in this subgroup.

We analyzed data from the Lipid Research Clinics (LRC) Coronary Primary Prevention Trial (CPPT)[7,8] to determine whether reported regular strenuous physical activity is related to the risk of coronary heart disease among asymptomatic hypercholesterolemic men.

Methods

Details regarding the design, implementation, and major results of the CPPT have been previously reported.[7,8] The CPPT was a double-blind, placebo-controlled clinical trial that tested the efficacy of lowering cholesterol levels with cholestyramine for the primary prevention of coronary heart disease (CHD). Twelve Lipid Research Clinics recruited 3,806 men, 35–59 years old, with primary hypercholesterolemia who were free of symptomatic CHD at entry. The men were required to have a plasma total cholesterol of 265 mg/dl (6.85 mmol/L) or greater (≥95th percentile) and plasma low-density lipoprotein cholesterol of 190 mg/dl (4.91 mmol/L) or greater at entry. Men with a plasma triglyceride greater than 300 mg/dl (33.07 mmol/L) or with type III hyperlipidemia were excluded.

Potential participants also were excluded from the trial if they had any of the following clinical manifestations of CHD:

- history of definite or suspect myocardial infarction;
- angina pectoris as determined by the Rose questionnaire;
- angina pectoris during submaximal exercise testing.
- various ECG abnormalities indicative of cardiovascular disease according to the Minnesota code: left bundle branch block, second or third degree heart block; two or more consecutive ventricular premature beats, left ventricular hypertrophy, R on T type ventricular premature beats, atrial flutter or fibrillation; or
- congestive heart failure.

Men with a positive exercise test result—ST segment depression or elevation of 1 mm or more—in the absence of other manifestations of CHD were not excluded from the CPPT.

Only men in good health and free of conditions associated with secondary hyperlipoproteinemia—diabetes mellitus, hypothyroidism, nephrotic syndrome, hepatic disease, hyperuricemia, and notable obesity—were included in the CPPT. Men also were excluded if they had moderate or severe hypertension—DBP ≥ 105 mmHg—or were receiving anti-hypertensive medication or had life-limiting or comorbid conditions such as cancer or non-atherosclerotic cardiovascular disease.

Participants were randomized into a treatment group that received cholestyramine and a control group that received a placebo; both groups were prescribed an identical moderate cholesterol lowering diet and attended Lipid Research Clinics every two months. Intervention by the clinic staff was restricted to the prescription of the study medication and the diet. The primary end-point for evaluating the treatment outcome in the CPPT was definite CHD death and/or definite non-fatal myocardial infarction.[7] All participants initially entered were followed up to the completion of the trial; the average length of follow-up was 7.4 years.

We restricted the present analysis to CPPT participants

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in the placebo plus diet group (1,899 men) who underwent graded exercise tolerance testing at entry (1,897 men). Because we wanted to minimize the likelihood that subclinical CHD might distort the relation between strenuous activity and CHD, we excluded participants with other resting electrocardiographic abnormalities according to the Minnesota code (24 men), such as minor Q wave or ST segment changes, a positive graded exercise tolerance test (266 men), or a history of intermittent claudication (three men). Because of their small number, 71 men who were not White were also excluded. Thus, 80 per cent (1,533 men) of the original CPPT placebo group was included in the analysis.

We determined the level of each participant's habitual physical activity by the response to two questions asked to select a target heart rate for the baseline graded exercise tolerance test. The first question was, "Do you regularly engage in strenuous exercise or hard physical labor?" Participants who answered no to the first question were classified as inactive. If the answer was yes, the participant was then asked, "Do you labor or exercise strenuously at least three times a week?" Since 77 per cent of the participants who answered yes to the first question reported strenuous exercise or hard physical labor at least three times a week, participants who responded yes to the first question were classified as active. Eighty-two per cent of those classified as inactive based upon the baseline assessment reported that they did not engage in regular strenuous activity at each annual follow-up examination in years 2 through 7 of the CPPT.

We examined the validity of self-reported physical activity indirectly by demonstrating expected relationships with other factors—resting heart rate, graded exercise tolerance test performance, daily dietary caloric intake, and subjective rating of activity level outside of work—measured at screening visit 2. In addition, prior analyses from both the LRC Prevalence Study and the CPPT had demonstrated that self-reported physical activity based upon the response to the same two questions was related to plasma levels of high-density lipoprotein cholesterol.[9,10]

Factors that might confound the relation between regular strenuous activity and CHD also were assessed at entry into the CPPT. These included: total cholesterol, low-density lipoprotein cholesterol (LDL-C) (pre-diet), high-density lipoprotein cholesterol (HDL-C) (pre-diet), age, family history of heart disease, cigarette smoking, systolic blood pressure, Quetelet index, alcohol consumption, fasting plasma glucose, education, and occupation.

We examined the incidence of CHD death and/or nonfatal myocardial infarction (MI) in relation to the self-report of regular strenuous physical activity at entry into the CPPT. We also examined the relation between regular strenuous activity and CHD in a proportional hazards model that adjusted for baseline differences between physically active and inactive men in other risk factors at entry into the CPPT.

Results

We indirectly validated our assessment of physical activity in our analysis sample by demonstrating expected relationships with other factors, shown in Table 1. As expected, on the average active males had a lower resting heart rate and a lower heart rate at the end of the second stage of the graded exercise tolerance test, a longer graded exercise tolerance test duration, and a higher daily dietary caloric intake, and rated their activity outside of work as somewhat more than average, compared to inactive men.

TABLE 1—Validation of Reported Regular Physical Activity

	Active (N = 489)	Inactive (N = 1044)
	Means	
Resting Heart Rate (beats/min)	75	80
Graded Exercise Test		
Stage 2 Heart Rate (beats/min)	129	137
Total duration (sec)	615	572
Daily Dietary Intake (cal/kg)	30	29
Activity Outside Work*	2	3

*Based upon a subjective rating on a scale from 1 to 5; 2 = somewhat more than average, 3 = average.

The baseline characteristics of the men included in our analysis are shown according to their self-reported activity status in Table 2. As expected, active males had higher levels of HDL-C, were less likely to smoke, and more likely to engage in occupations that required manual labor, compared to inactive males. In addition, active males consumed more alcohol and were slightly more likely to have a family history of premature CHD. Total cholesterol, LDL-C, age, systolic blood pressure, Quetelet index, fasting glucose, and the proportion of college graduates were similar in the two groups.

The incidence of coronary heart disease death and/or nonfatal infarction was similar among active and inactive males (Table 3). Thus, physical activity was not related to the risk of CHD as reflected by our estimates of relative risk. Adjustment by the proportional hazards model for differences in the baseline characteristics and the length of follow-up did not alter these findings. In addition, our findings

TABLE 2—Baseline Characteristics According to Reported Regular Physical Activity

Characteristics	Active (N = 489)	Inactive (N = 1044)
	Means	
Total cholesterol (mg/dl)†	292	291
LDL cholesterol* (mg/dl)†	219	218
HDL cholesterol* (mg/dl)†	47	44
Age (years)	47	47
Systolic BP (mm Hg)	123	123
Quetelet index (gr/sq cm)	2.63	2.63
Fasting plasma glucose (mg/dl)‡	98	98
Alcohol consumption (g/day)	10.9	9.5
	Percentage	
Current smokers	31	39
Family history heart disease	27	24
College graduate	39	37
Occupation		
Manager	35	37
Clerical	36	43
Manual	29	19

*Pre-dietary intervention
†Conversion factor for mmol/L: 0.02586
‡Conversion factor for mmol/L: 0.05551

TABLE 3—Unadjusted Incidence of Coronary Heart Disease According to Reported Regular Physical Activity

Outcome	Active (N = 489)	Inactive (N = 1044)	Relative Risk (95% CI)
	Cases/1000 persons/year(N)		
Definite CHD death and/or definite nonfatal myocardial infarction	11.9 (43)*	12.6 (98)*	.94† (.68, 1.38)
Definite CHD death	1.9 (7)	1.9 (15)	
Definite nonfatal infarction	10.2 (37)	11.1 (85)	
Definite or suspect CHD death and/or nonfatal myocardial infarction	17 (59)*	16.8 (129)*	1.01 (.75, 1.37)
Definite or suspect CHD death	2.4 (9)	2.2 (18)	
Definite or suspect nonfatal myocardial infarction	14.6 (52)	15.0 (115)	

*A man experiencing a myocardial infarction and CHD death is counted once in this category; hence, this line is not the sum of the following two lines.

†After adjustment for age, smoking, LDL-cholesterol, family history of CHD, and occupation through use of a proportional hazards model, the relative risk = .98, 95% confidence interval = .68, 1.41.

were similar when we considered only men who reported strenuous activity at least three times a week as active.*

Examination of survival curves for combined fatal CHD and non-fatal MI according to baseline activity status confirmed the lack of an effect of reported activity on the risk of CHD during the entire follow-up period.* These findings were similar for fatal CHD and non-fatal myocardial infarction examined separately.*

Discussion

Our data suggest that reported regular strenuous activity or hard physical labor is not related to the incidence of CHD death and/or non-fatal myocardial infarction in asymptomatic hypercholesterolemic men. Although this finding may reflect a lack of effect, there are several alternative explanations that need to be considered.

First, misclassification of activity status might have obscured an effect of regular strenuous activity on the risk of CHD. We used a self-report of regular strenuous exercise or hard physical labor to characterize the level of habitual physical activity. The intensity, frequency, and duration of participation in specific types of physical activity were not measured. However, 77 per cent of the men in our analysis sample who reported some strenuous activity indicated that they regularly engaged in that activity for three or more times per week. In addition, we indirectly demonstrated the validity of self-reported activity in our study population. It is unlikely that the differences observed between active and inactive men in resting and graded exercise stage 2 heart rate, treadmill test duration, and HDL-C occurred by chance alone ($P < .001$ for each comparison). However, our analysis does not preclude the possibility of some misclassification, as suggested by the absence of larger differences between active and inactive men in these measures that indirectly reflect variability in strenuous physical activity. Nevertheless, we consider the self-report of regular strenuous activity or hard physical labor an indication that the man was physically active and not sedentary at entry into the CPPT.

Second, the statistical power of our analysis to detect an association between regular activity and CHD was somewhat limited, as reflected by the width of the 95% confidence limits for our estimates of relative risk. It is possible that we missed a decrease or increase in risk associated with regular activity of approximately 30 per cent, but it is unlikely that we missed the 50 per cent risk reduction demonstrated in previous epidemiologic studies of heterogeneous populations.[1–4] Furthermore, our best estimate of effect—the point estimate—suggests a very weak or no association between reported activity and CHD in our study population.

Third, males who were older, hypertensive, obese, or diabetic were excluded from the CPPT. Since we have previously suggested that males who are older, hypertensive, or obese might particularly benefit from activity,[11] these selection criteria might have influenced our estimate of effect. Nevertheless, we consider the presence of hypercholesterolemia, defined in the CPPT as a total cholesterol ≥ 95th percentile for middle-aged males, *the* major difference between our study population and those of previous studies.[1–4] For these reasons, our results are not directly comparable to those from other study populations, and it might not be appropriate to generalize our finding of no effect to persons with lower levels of plasma cholesterol and/or other CHD risk factors.

Regular strenuous activity has been suggested as an adjunct to the treatment of hypercholesterolemia.[12] In part, this recommendation is based upon the relationship between activity and HDL-C.[5] Active men had a higher plasma level of HDL-C at entry into the CPPT. The difference in HDL-C between active and inactive men was greater than that observed in exercisers versus controls in a one-year randomized study of the effects of exercise on plasma lipoproteins[13] and in a recent report from the Multiple Risk Factor Intervention Trial where leisure-time physical activity was inversely related to the incidence of first major CHD events and overall mortality.[14] Based upon the relation between HDL-C level and CHD in the CPPT placebo group, we might have expected active men to have an overall 10 per cent reduction in CHD.[15] Our analysis did not have adequate statistical power to detect an effect of this magnitude. However, it is also possible that the potential benefit of activity is balanced by the cardiac risks of activity in hypercholesterolemic men.

Our data suggest that reported regular strenuous physical activity may not be related to the risk of CHD among asymptomatic hypercholesterolemic middle-aged men. Thus, regular activity might not mitigate the effect of life-long hypercholesterolemia. For now, greater attention to other potential interventions that reduce the risk of CHD among hypercholesterolemic men[7] seems appropriate.

*Data available from authors on request.

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'Hospitals, Health Care Professionals and AIDS' Is Conference Theme

The American Society of Law & Medicine (ASLM), Public Responsibility in Medicine and Research (PRIM&R), and the Harvard AIDS Institute will hold a major national conference on "Hospitals, Health Care Professionals and AIDS" at the Boston Park Plaza Hotel, December 1-2, 1988.

The meeting will focus on such charged and troubling issues as: the ethical duty to care for AIDS patients; discrimination against health care providers and patients; financing care and reimbursement; research and clinical initiatives; the rights of AIDS patients to experimental treatment; infection control; screening of patients; HIV-infected surgeons.

At the conference, issues will be discussed by leading national speakers from the fields of clinical research, medicine, law, support groups for AIDS patients and bioethicists. Conference registrants will also have an opportunity to participate in topical workshops and plenary sessions during the two-day meeting.

For further conference and registration information, contact Robyn Carey, Assistant Director, PRIM&R, 132 Boylston Street, Boston, MA 02116. Tel: (617) 423-4112; or Sharon Paaso, Associate Director, ASLM, 765 Commonwealth Avenue, Boston, MA 02215. Tel: (617) 262-4990.

The Persistence of *Shigella flexneri* in the United States: Increasing Role of Adult Males

ROBERT V. TAUXE, MD, MPH, ROBERT C. MCDONALD, MD, NANCY HARGRETT-BEAN, PHD, AND PAUL A. BLAKE, MD, MPH

Abstract: The annual reported isolation rate of *Shigella flexneri* decreased from 1964 to 1973, but has remained constant since then at 1 per 100,000. Between 1975 and 1985, the median age of males from whom *S. flexneri* was isolated rose from 5 to 26 years. During this time, the isolation rate of *S. flexneri* rose more than five-fold among men, did not change in adult women, and decreased in children. By 1985, 23 per cent of reported *S. flexneri* isolates came from men aged 20–49. Increased male homosexual transmission of *S. flexneri* is a possible explanation for these findings. (*Am J Public Health* 1988; 78:1432–1435.)

Introduction

Shigella flexneri has been replaced by *S. sonnei* as the most frequently reported *Shigella* serotype in many industrialized nations.[1] In the United States, the number of reported *S. flexneri* isolates declined steadily from 1964 to 1973, and the proportion of *Shigella* isolates that were *S. flexneri* dropped from 60.6 per cent to 15.5 per cent.[2] This decline stopped in 1974, and the number of reported isolates of *S. flexneri* has remained nearly constant since. We reviewed national *Shigella* surveillance data from 1967 to 1985 to examine and explain the persistence of *S. flexneri* in the United States.

Methods

The national *Shigella* surveillance system, organized in 1963, reached full national participation in 1967; since then all states except California have contributed continuously.* It is a passive, laboratory-based surveillance system which receives weekly reports of isolates and accompanying demographic information from state health departments. Duplicate isolates are excluded. No clinical information is provided, and asymptomatic excretors are not distinguished from the acutely ill. It has been estimated that 5 per cent of *Shigella* infections are reported.[3] Reporting forms, coding and data management have remained constant since the inception of the system.

Because the specific serologic subtype is often not reported, we grouped serotypes as *S. flexneri* 1, (serotypes 1a, 1b, and 1 unspecified); *S. flexneri* 2, (2a, 2b, and 2 unspecified); *S. flexneri* 3, (3a, 3b, 3c, and 3 unspecified); and *S. flexneri* 4, (4a, 4b, and 4 unspecified). Age-specific isolation rates were calculated based on the annual 49-state US civilian population. City-specific data were derived from the county-based *Shigella* surveillance system, by using the counties included in the Standard Metropolitan Statistical Areas (SMSAs) of the 21 largest cities outside of California.[4]

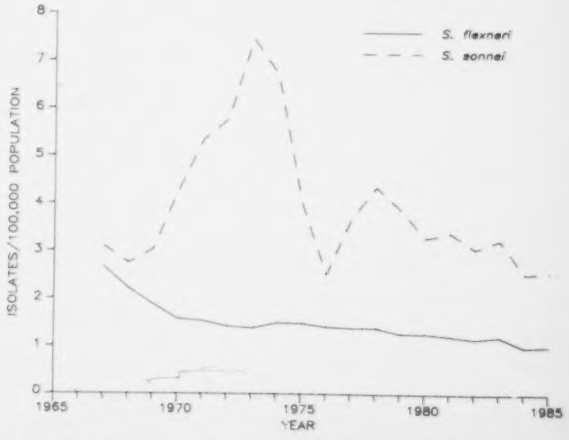
Address reprint requests to Robert V. Tauxe, MD, MPH, Chief, Epidemiology Section, Enteric Diseases Branch, Division of Bacterial Diseases, Center for Infectious Diseases, 1-54 28, CDC, Atlanta, GA 30333. Dr. Hargrett-Bean and Dr. Blake are also with that division at CDC. Dr. McDonald, is an epidemiology elective student with the CDC Epidemiology Program Office, presently at the Department of Medicine, Washington University, St. Louis, MO. This paper, submitted to the Journal February 12, 1987, was revised and accepted for publication April 4, 1988.

*California did not report isolates between 1968 and 1973, and has not reported age-, sex-, or residence-specific data since then.

Results

The isolation rate of *Shigella flexneri* declined through 1973, and has been relatively stable since then at approximately 1 per 100,000 population (Figure 1). In the past, persons in custodial institutions and persons on Indian reservations have been at high risk for *S. flexneri* infections. However, the number of *S. flexneri* isolates reported from these groups has been decreasing, particularly from custodial institutions (Table 1). Although many isolates do not have a reported residence, these two groups accounted for only 5.9 per cent of all *S. flexneri* isolates for which a residence was reported during 1979–84.

A profound demographic shift in persons from whom *S. flexneri* isolates were reported began in 1976 (Figure 2). The reported age began to increase, particularly in males, whose median age rose from 5 years in 1975 to 26 years in 1985. This is the combined effect of a decrease in the age-specific isolation rates for children and an increase in the isolation rate among adult men (Figure 3). The greatest increase occurred among men 30 to 39 years of age. In this age group the isolation rate of *S. flexneri* climbed fom 0.26 per 100,000 in 1970 to 1.41 per 100,000 in 1985 (Figure 4). By comparison the isolation rate among women in the same age group did not change, and the rate among young children decreased substantially. The increase in median age among females since 1983 is the result of the decreased isolation rate among



*REPORTED TO THE NATIONAL SHIGELLA SURVEILLANCE SYSTEM, EXCLUDING CALIFORNIA.

FIGURE 1—Isolation Rate of *S. flexneri* and *S. sonnei* Reported through the National *Shigella* Surveillance System, 49 States, United States, 1967–85

TABLE 1—*S. flexneri* Isolates Reported from High-Risk Groups in the United States, by 6-Year Interval, 1967–84

Interval	No. *S. flexneri* with Residence Reported	Mental Institutions # Isolates	Mental Institutions (% of Total)	Indian Reservations # Isolates	Indian Reservations (% of Total)
1967–72	8257	3136	(38.0)	455	(5.5)
1973–78	6588	1029	(15.6)	515	(7.8)
1979–84	5726	72	(1.3)	264	(4.6)
1967–84	20571*	4237	(20.6)	1234	(6.0)

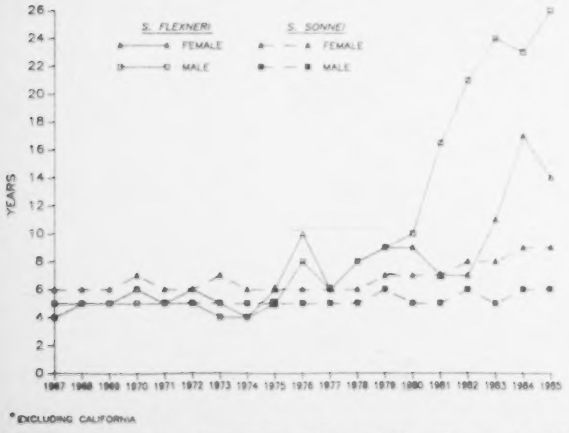
*Representing 29 per cent of all reported *S. flexneri*.

children, and not to an increase in the isolation rate among adult women. In 1970, 5.5 per cent of all reported *S. flexneri* isolates came from men aged 20–49; this proportion increased to 23 per cent in 1985.

These changes in age- and sex-specific isolation rates can be seen in the individual serologic subgroups of *S. flexneri*. The isolation rate of serogroup 1 among men increased sharply in 1980, without a similar increase that year among adult women (Figure 5). The other *S. flexneri* serogroups also show striking temporary increases among adult men without parallel increases among women or children: *S. flexneri* four in 1974, *S. flexneri* two in 1978, *S. flexneri* three again in 1983, and *S. flexneri* four in 1984.

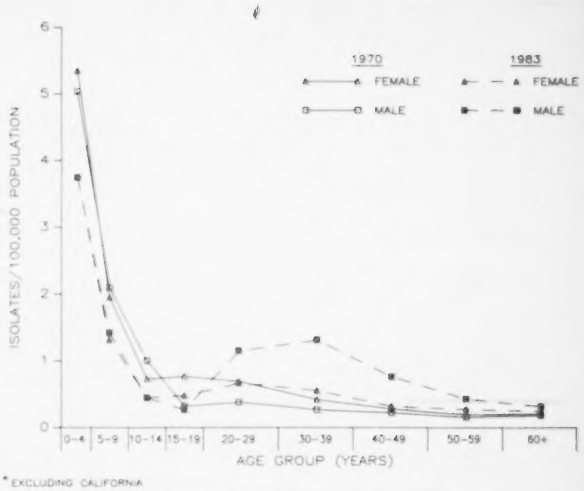
The increase in *S. flexneri* isolations among adult men is particularly evident in large cities (Table 2). In the 21 largest cities in the United States outside of California, in the period 1967–72, isolates of *S. flexneri* from women 20–49 years of age equaled or outnumbered those from men of the same age in all but two of these cities. In the period 1973–78, isolates from men became more frequent, exceeded those from women in eight of these cities, and accounted for over half of all *S. flexneri* reported in the US in that age- and sex-group. In the period 1979–84, these trends continued, and isolates from men exceeded those from women in 16 of these cities.

In contrast, the median age of persons with *S. sonnei* has not shifted markedly (Figure 2). In 1970, men aged 20–49 years accounted for 8 per cent of isolates of *S. sonnei*, in 1985 this proportion was 10 per cent. In 17 of the 21 cities



*EXCLUDING CALIFORNIA

FIGURE 2—Median Age of Persons from Whom Isolation of *S. flexneri* and *S. sonnei* Was Reported, by Sex, 49 States, United States 1967–85



*EXCLUDING CALIFORNIA

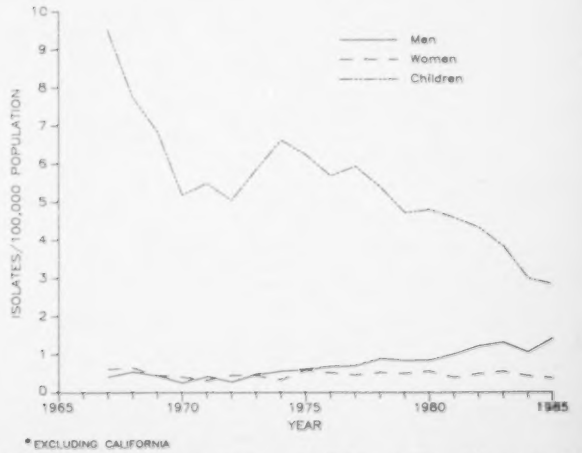
FIGURE 3—Age- and Sex-specific Isolation Rates of *S. flexneri*, 49 States, United States, 1970 and 1983

examined above, *S. sonnei* was isolated from more women than men aged 20–49 in the six-year period 1979–84. Sixty-three per cent of *S. dysenteriae* from persons ages 20 to 49 in 1967–72 came from men, 46 per cent in 1973–78, and only 44 per cent in 1979–84.

Discussion

Interpretation of national *Shigella* surveillance data is hampered by the limited information reported with each isolate, by regional variation in case ascertainment and reporting, and by changes in the delivery of medical care over time. However, laboratory methods for culturing and serotyping *Shigella* have changed little in the last 20 years, and the surveillance system itself has remained virtually constant since its inception. Internal comparisons suggest secular trends for the nation as a whole.

S. flexneri has persisted in the United States despite a steady decline in the rate of isolation from children, and a



*EXCLUDING CALIFORNIA

FIGURE 4—Isolation Rates of *S. flexneri* in Children less than 5 Years old and Men and Women Aged 30–39, 49 States, United States, 1967–85

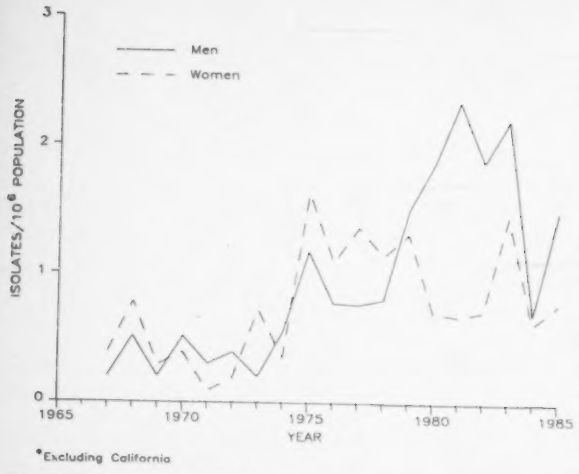


FIGURE 5—Isolation Rate of *S. flexneri* Serogroup 1 in Men and Women Aged 30–39, 49 States, United States, 1967–85

decrease in the number of reported isolates from Indian reservations, and custodial institutions. This persistence has been accompanied by a striking increase in the isolation rate among adult men; approximately one-fourth of reported *S. flexneri* isolates now come from adult men. In many large cities, infection restricted to adult men appears to have supplanted the traditional cycle of *S. flexneri* infections in children and their caretakers. Individual serogroups of *S. flexneri* have gained national predominance among adult men in a sequential fashion without similar increase in women or children. These observations suggest that the urban-dwelling man aged 20–49 is at particular risk for *S. flexneri*, and that transmission restricted to this group may explain much of the persistence of the serogroup.

There are several possible explanations for the increasing role of the adult male. Men may participate more in the care of ill children, but this would not explain selective increases in serotypes restricted to adult men. Men may travel more than women, although one might expect travel-associated shigellosis to represent a broad spectrum of serotypes, rather than successive waves of individual serotypes, and *S. dysenteriae*, a travel-associated serotype, is more frequent among women than men. An increase in the number of homeless men, particularly those discharged from mental institutions, may play a role, although *S. flexneri* has not been documented as a particular problem of the homeless. An increase in adult male immigration from Latin America to the urban United States is a likely but undocumented contributing factor. Male homosexual transmission is also likely to be an important factor, if *S. flexneri*, like other established homosexually transmitted diseases, has become common among young, urban male homosexuals.[5] The relative importance of these explanations cannot be ascertained from the national surveillance data.

Male homosexual shigellosis was first noted in San Francisco in 1974, the year that the national upward trend in *S. flexneri* isolates from men began.[6] In that city, *S. flexneri* 2a was initially predominant among male homosexuals, followed by *S. flexneri* 3a, and then *S. sonnei*.[7] In Seattle the following year, *S. flexneri* 3 was the predominant serogroup among male homosexuals and *S. sonnei* also appeared to be transmitted among male homosexuals; 30 per cent of all shigellosis reported in 1975 occurred in male homosexuals.[8] Among men with shigellosis at the New York Hospital in 1970–75, 58 per cent of those without recent travel were homosexual, compared to 10 per cent of those with recent travel.[9] By 1985, shigellosis in San Francisco was largely a disease of homosexual men: 82 per cent of all reported *Shigella* was *S. flexneri*, and 66 per cent of patients with shigellosis were homosexual or bisexual men.[10] The relative rarity of *S. sonnei* in San Francisco is curious, since this is by far the most common serogroup in the US, and raises the question of whether *S. flexneri* is more efficiently transmitted than *S. sonnei* by homosexuals. Although immunosuppressed adults may develop severe *Shigella* bacteremia, the effect of AIDS (acquired immunodeficiency syndrome) on *Shigella* transmission is not known.[11]

It remains to be established whether homosexual transmission explains the nationwide increase in *S. flexneri* isolations from adult men. If it does, reducing the incidence of these infections may require new control strategies, beyond emphasizing handwashing and food hygiene. The homosexual practices likely to transmit *Shigella* are oral-anal and oral-genital sex. Changes in the prevalence of these practices because of concern about AIDS could have some impact on *S. flexneri*.[6,8,12–14] In San Francisco, reported shigellosis decreased by 30 per cent from 1981 to 1985.[10] This decrease may be analogous to recent reports of decreases in rectopharyngeal gonorrhea in men in New York City,[15] and in gonorrhea in men in Denver,[16] and London.[17] However, the continued nationwide increase in *S. flexneri* isolations from adult men through 1985 suggests that such behavior changes are not widespread, or that non-sexual transmission among adult men is also important. In some areas, case-investigation of *S. flexneri* reported from men may help identify those in need of health education about AIDS.

TABLE 2—Reported *S. flexneri* Isolates from 20- to 49-year-old Persons by Sex in the 21 Largest United States Cities,* by Year of Report, 1967–84

Population Group	1967–72 Male/Female	1973–78 Male/Female	1979–84 Male/Female
21 Cities Total:	187/301	721/499	1240/625
(Proportion male)	(0.38)	(0.59)	(0.66)
US total for age group*	878/1192	1402/1351	2383/1474
(Proportion male)	(0.42)	(0.51)	(0.62)
Proportion of US total for age and sex group accounted for by the 21 cities	0.21/0.25	0.51/0.37	0.52/0.42

*Excluding California

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OTA Special Report: Biology, Medicine, and the Bill of Rights

By effectively forcing people to make life and death decisions about themselves, about other people, and even future generations, the new genetic and medical technologies are raising questions about the meaning and scope of constitutional principles, according to a recent report from the Congressional Office of Technology Assessment (OTA).

Constitutionally, the central question becomes: would State regulation of these decisions impinge on some individual liberty that is guaranteed by the Bill of Rights? And, if so, is the individual's interest in exercising that right outweighed by the contrary interest of the State — considered to be the "public interest"?

Whether the government can or should regulate biotechnology is not an abstract question, according to Representative Robert W. Kastenmeier (D-WI) of the House Committee on the Judiciary. "It involves matters of birth and death, heredity and new life forms, disease and aging," he noted. In view of these and similar policy issues, he requested OTA to assess how the laws, institutions and scientific research might move forward together. This report is the result of that effort.

According to OTA, other decisions include the patenting of living organisms, using fetal tissues for research, human cell therapy, prenatal diagnosis and treatment, mandatory testing for genetic or infectious diseases, the right to treatment, and the right to refuse treatment. Is there a constitutional right to do research? Should there be areas of "forbidden knowledge"? What values should be reflected in federal research funding allocation and federal guidelines?

Some traditional public health practices — constitutionally long- established under State police powers — are almost certain to be challenged anew because of today's broader interpretation of individual rights of privacy and autonomy, according to OTA, This is occurring, for example, in the context of the AIDS epidemic with regard to mandatory reporting, contact tracing, mandatory testing, occupational or educational restrictions, and quarantine.

Copies of the OTA 124-page special report, *Biology, Medicine, and the Bill of Rights,* are available from the Government Printing Office, Superintendent of Documents, Washington, DC 20402. Tel: (202) 783-3238. GPO stock number 052-003-01133-6. The price is $4.25.

Improving the Life-Course Development of Socially Disadvantaged Mothers: A Randomized Trial of Nurse Home Visitation

David L. Olds, PhD, Charles R. Henderson, Jr., Robert Tatelbaum, MD, and Robert Chamberlin, MD

Abstract: We evaluated a comprehensive program of prenatal and postpartum nurse home visitation for socially disadvantaged women bearing first children. Eighty-five per cent of the participating women were either teenagers (<19 years at registration), unmarried, or of low socioeconomic status. Women were randomly assigned to either nurse home visitation or comparison services (free transportation for prenatal and well-child care and/or sensory and developmental screening for the child). During the first four years after delivery of their first child, in contrast to their counterparts in the comparison group, nurse-visited White women who had not graduated from high school when they registered in the study returned to school more rapidly; nurse-visited, poor, unmarried White women showed an 82 per cent increase in the number of months they were employed, had 43 per cent fewer subsequent pregnancies, and postponed the birth of second children an average of 12 months longer. (*Am J Public Health* 1988; 78:1436–1445.)

Introduction

In 1985 there were nearly 500,000 births to women under 20 years of age in the United States and over 800,000 to women who were unmarried, regardless of their age (12.7 per cent and 22.0 per cent of all births, respectively).[1] Adolescent and out-of-wedlock parenthood create substantial risks for both mothers and children. Adolescent mothers are in jeopardy of dropping out of school, having rapid successive pregnancies, and becoming dependent on welfare.[2–6] Children of adolescent mothers are more likely to be afflicted with a variety of health and developmental problems.[7–13] These maternal and child dysfunctions are even more likely to occur if an adolescent mother gives birth to a second child soon after the first.[4,5,14] Out-of-wedlock births create economic hardships and corresponding pressure on women to take low-paying jobs and/or to rely on Aid to Families with Dependent Children (AFDC).[15] Children born out of wedlock are also at risk for a host of problems,[16–19] but the extent to which those problems are due to the mother's single-parent status as opposed to poverty or young age is unclear.[15,20] Government expenditures on the sequelae of adolescent and out-of-wedlock childbearing are enormous.[21,22]

During the past 20 years, numerous programs have been developed either to prevent unintended pregnancy in these socially disadvantaged groups[23,24] or to reduce its consequences.[25–33] The effectiveness of many of these programs is obscurred, however, because of flaws with the evaluation designs.[34–36] Furthermore, many services for these groups fail to address the full array of stressful conditions with which socially disadvantaged parents must contend as they try to improve their life situations.

Even though it is common for visiting nurses to help new parents reduce unwanted subsequent pregnancies, return to school, and find employment, there are no studies that have examined the effectiveness of nurse home visitation as a means of improving maternal life-course development. The present study was established to test a comprehensive and intensive program of prenatal and postnatal nurse home visitation aimed at improving the outcomes of pregnancy, early childrearing, and life-course development of women who were either teenaged, unmarried, or poor, and bearing first children.

The program was based on the premise that nurse home visitors are in an optimal position to identify and help change factors in the family environment that interfere with maternal health habits, infant caregiving, and personal accomplishments in the areas of work, education, and family planning. In earlier papers we reported that the program improved the conditions and outcomes of pregnancy and reduced the incidence of caregiving dysfunction; many of the positive outcomes of the program were concentrated in families at highest risk for particular problems in these areas.[37,38] The intervention also was designed explicitly to promote maternal educational and occupational achievements and to reduce the number of subsequent unintended pregnancies. This paper evaluates the effectiveness of the program in these areas.

Methods

Research Design

The study design consisted of a randomized clinical trial. Participating families were assigned at random (details in Appendix) to one of the four treatment groups outlined in Table 1.

Treatment Conditions

Treatment 1—When the children were one and two years of age, an infant specialist hired by the research project screened them for sensory and developmental problems and referred those with suspected problems to other specialists

TABLE 1—Services Provided in Each of the Four Treatment Groups

Services Provided	Treatment Group 1 N = 90	2 N = 94	3 N = 100	4 N = 116
Sensory and developmental screening at the 12th and 24th months of life	+	+	+	+
Free transportation to regular prenatal and well-child visits		+	+	+
Nurse home visitation during pregnancy			+	+
Nurse home visitation during the first 2 years of life				+

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for further evaluation and treatment. No services were provided through the research project during pregnancy.

Treatment 2—Families were provided free transportation for regular prenatal and well-child care at local clinics and physicians' offices through a contract with a local taxicab company, as well as the sensory and developmental screening outlined in Treatment 1.

Treatment 3—Families were provided a nurse home visitor during pregnancy, in addition to the screening and transportation services. The nurses visited families approximately once every two weeks and made an average of nine visits during pregnancy. Each visit lasted approximately 1 hour and 15 minutes.

Treatment 4—Families received the same services as those in Treatment 3, but in addition the nurse continued to visit until the children were two years of age. For six weeks following delivery the nurses visited families every week; from six weeks to four months they visited every three weeks; from 14 to 20 months they visited every four weeks; and from 20 to 24 months they visited every six weeks. Under predetermined crisis conditions, the nurses visited weekly. As during pregnancy, the visits lasted approximately 1 hour and 15 minutes.

Program Focus on Maternal Life-Course

Beginning during pregnancy, the nurses attempted to form an effective therapeutic relationship with the women by emphasizing the women's personal strengths. For example, the nurses:

- encouraged the women to clarify plans for completing their education, returning to work, and bearing additional children;
- stressed that the decision to return to school or seek employment after delivery should be made after fully considering what was in the women's own and their babies' best interests;
- helped interested women find appropriate educational and vocational training services and make concrete plans for child care;
- advised them in finding jobs and interviewing;
- showed the women and their partners birth control devices and discussed the advantages of different methods of family planning.

The discussions of family planning were carried out in the context of the women's desires for continued education, work, and achieving what they considered to be their optimal family size. When the women so desired, their husbands, boy friends, mothers, and other family members were encouraged to participate in the home visits. Families were also referred to other health and human services as needed.

The nurses used a detailed curriculum to guide their activities, but tailored the specific content of their home visits to the individual needs of each family. They refrained from imposing family planning methods on couples who had already decided that they wanted to have large families and their children spaced closely together.

Five registered nurses (none with baccalaureate degrees) were hired through a nonprofit private agency expressly for this experimental program. Before being assigned families in the experimental program, they participated in a three-month training program in which they each began working with two "pilot" families. Each nurse followed a caseload of 20–25 families from pregnancy through the child's second birthday. The nurses worked in two-person teams, with each member serving as a backup for her partner. By visiting one another's families at least once during both the pregnancy and the postpartum period, the members of the team were able to serve as ongoing supports for one another on a day-to-day basis. Detailed descriptions of the program are provided elsewhere.[39–42]

The Setting

The study was carried out in a small, semirural county of approximately 100,000 residents in the Appalachian region of New York State. At the time the study was initiated, the community was well served from the standpoint of both health and human services. In spite of this abundance of services, the community has consistently exhibited the highest rates of reported and confirmed cases of child abuse and neglect in the state.[43] Moreover, between 1967 and 1982 the number of manufacturing jobs in the area dropped from 17,400 to 9,000 (a 48 per cent decline)[44]; in 1981, the community was rated the worst Standard Metropolitan Statistical Area in the country in terms of its economic conditions.[45]

The Sample

Women were actively recruited if, at intake, they had no previous live births and had any one of the following characteristics that predispose to infant health and developmental problems: young age (<19 years); single-parent status; low socioeconomic status (SES). The study design nevertheless allowed the enrollment of any woman who asked to participate and who was bearing her first child. This avoided creating a program that was stigmatized as being exclusively for the poor; furthermore, by creating sample heterogeneity, it enabled us to determine whether the effects of the program were greater for families at higher risk. All women were enrolled before their 30th week of pregnancy. They were recruited through a free antepartum clinic sponsored by the health department, the offices of private obstetricians, Planned Parenthood, the public schools, and a variety of other health and human service agencies. Approximately 10 per cent of the target population was missed because of late registration for prenatal care. An additional 10 per cent was missed because some eligible women from the offices of private obstetricians were not referred. Between April 1978 and September 1980, 500 women were interviewed and 400 were enrolled. There were no differences in age, marital status, or education between those women who participated and those who declined. Eighty-eight of 442 Whites (20 per cent) declined to participate compared to two of 48 others (mostly Black). At registration, 47 per cent of the participating women were under 19 years of age, 62 per cent were unmarried, and 61 per cent came from families in Hollingshead's social classes IV and V (semiskilled and unskilled laborers).* Fifteen per cent of the women were not at risk according to the age, marital status, or SES criteria, and 23 per cent possessed all three risk characteristics.

The 46 non-Whites were excluded from the analyses reported here because this sample was too small to provide adequate subclass sizes when other factors of importance were included in the statistical model. The analysis of the non-White sample is presented elsewhere.[46]

*Hollingshead A: Four factor index of social status (unpublished manuscript), Yale University, 1976. Hollingshead's index was adapted slightly to accommodate the variety of household compositions found in our sample. Details of the procedure can be obtained from the first author.

Data Collection

The interviews that produced data for this report were carried out at registration (prior to the 30th week of pregnancy) and at the 6th, 10th, 22nd, and 46th months of the children's lives. Except in a small number of cases where women inadvertently disclosed their treatment assignments, all interviews were conducted by staff members unaware of the women's treatment.

Records from the county department of social services were reviewed by a social services employee (who also was unaware of families' treatment assignment) to determine the number of days that the women and their children received public assistance during the period from the index child's birth to fourth birthday. This review was carried out on women who were either unmarried, teenagers, or from social classes IV or V, for these conditions made them more likely to use public assistance; or whose interview data concerning public assistance use indicated that they had been recipients. Those cases in which women were believed to receive public assistance were treated as missing data if they moved out of the county for more than two months, even if partial public assistance data (based on the record review) were available. There were two exceptions to this rule:

- If the results of the record review indicated that the woman had been on public assistance for fewer days than indicated by her interview data, and she had moved out of the county during the interval under consideration, the woman's interview was used as the source of data.
- If the woman had temporarily moved out of the county and there was no indication of public assistance use before or after the move (from either the interview or record review), then 0 days on public assistance was entered as the outcome of our assessment.

Combining Treatments 1 and 2 for Analysis

Treatments 1 and 2 were combined for purposes of analysis after it was determined that there were no differences between these two groups in their use of routine prenatal and well-child care, the primary means by which transportation was hypothesized to affect outcomes. The combination of Treatments 1 and 2 is referred to hereafter as the comparison group. Treatment 4 is referred to as the nurse-visited (NV) group, or in the tables as "NV Pregnancy/Infancy." Treatment 3 is labeled in the tables as "NV Pregnancy" (nurse-visited during pregnancy). Estimates of treatment differences focused on the contrast of NV-Pregnancy/Infancy vs the comparison group for both the whole sample and those groups defined at risk. The statistical model and methods of analysis are discussed in the Appendix.

Results

Preintervention Treatment Conditions

Table 2 shows that before assignment to treatment conditions, the nurse-visited and comparison-group women were equivalent on all standard sociodemographic background characteristics. They differed, however, on some social support and psychological characteristics: In contrast to the women in the comparison group, there was a trend for the nurse-visited group to expect less accompaniment to labor and delivery; the nurse-visited unmarried women had a greater sense of control; and the poor, unmarried teenagers assigned a nurse reported receiving greater support from their boy friends. Because sense of control and husband/boy friend support were more consistently related to the outcomes of this study than was expected accompaniment to labor and delivery, the potential bias created by this initial nonequivalence was handled by including the first two variables in the statistical model as covariates.

Attrition

During the first four years after delivery, the women's rates of attrition varied from 15 per cent to 21 per cent (depending on assessment period) and there were no differences across treatments in the proportion of subjects with completed assessments. The nurse-visited women who discontinued the program, however, tended to have a greater sense of control over their lives than did those in the control group who discontinued. An examination of the reasons for these women's dropping out of the program indicated that they had either moved or miscarried. Because a larger number of women with greater sense of control discontinued in the nurse-visited group than did those in the comparison group, the preintervention treatment difference in sense of control for unmarried women was reduced in the sample on which 46th-month interviews were carried out.

Educational Achievement

Table 3 shows that among women who had not graduated from high school when they registered for the study, 59 per cent of the nurse-visited and 27 per cent of the comparison group had either graduated or had enrolled in an educational program by their 6th month postpartum. By the 10th month postpartum, the effect of the program held only for those women who had been unmarried at registration. There were no treatment differences in the proportion of women who graduated or remained in high school at the 22nd month postpartum, and there were no treatment differences by the 46th month postpartum in overall educational achievement.

Employment, Child Care, and Public Assistance

Table 4 shows that between birth and the 22nd month after delivery, nurse-visited, poor, unmarried older women had worked two and one-half times longer than their counterparts in the comparison group. By the 46th month postpartum, the nurse-visited, poor unmarried teenagers also had begun to work more than their counterparts in the comparison group, leading to an 82 per cent increase in the number of months worked by nurse-visited, poor, unmarried women (both teenagers and older) in contrast to poor, unmarried women in the control group. A qualitative analysis of the employment data revealed that at the time of the 46th month interview, most of the women who worked held unskilled labor and service positions; some held semiskilled jobs; a few were clerical and sales workers.

Table 4 also indicates that at both the 10th and 22nd month interviews, the nurse-visited, poor, unmarried teenagers reported greater concern about eventually finding employment than did the poor, unmarried teens in the comparison group, in spite of equivalent employment histories at those points in time. Among the poor, unmarried older women, those who were nurse-visited reported at the 22nd month interview less concern about finding employment.

At the 10th month postpartum, in contrast to their comparison-group counterparts, the nurse-visited, poor, unmarried older women reported that other family members helped them more with child care. During the first two years after delivery, the nurse-visited, poor, unmarried older women were on public assistance 157 fewer days than poor, unmarried older women in the comparison group (a 40 per

TABLE 2—Estimates of Preintervention Treatment Differences and 95% Confidence Intervals for Maternal Background Characteristics

Dependent Variable	Treatment Group	Whole Sample		Poor, Unmarried Women	
		$\bar{X}$	N	$\bar{X}$	N
Hollingshead social class score (range: 22–128)[a]	Comparison	58.02	(165)	41.05	(67)
	NV Pregnancy	57.59	(90)	41.63	(36)
	NV Pregnancy/Infancy	59.21	(99)	43.95	(37)
	Comparison-NV Infancy	−1.19		−2.90	
	C.I.	(−6.70, 4.32)		(−7.48, 1.68)	
Maternal age at registration (range: 14–34)	Comparison	19.28	(165)	18.79	(67)
	NV Pregnancy	19.52	(90)	19.08	(36)
	NV Pregnancy/Infancy	19.50	(99)	18.90	(37)
	Comparison-NV Infancy	−0.22		−0.11	
	C.I.	(−0.82, 0.38)		(−0.99, 0.77)	
Maternal education (years completed) (range: 7–17)	Comparison	11.25	(165)	10.60	(67)
	NV Pregnancy	11.58	(90)	10.93	(36)
	NV Pregnancy/Infancy	11.32	(99)	10.68	(37)
	Comparison-NV Infancy	−0.07		−0.08	
	C.I.	(−0.41, 0.27)		(−0.57, 0.41)	
Maternal sense of control[b] (range: 7–14)	Comparison	12.21	(165)	11.95	(67)
	NV Pregnancy	12.31	(90)	12.30	(36)
	NV Pregnancy/Infancy	12.44	(99)	12.33	(37)
	Comparison-NV Infancy	−0.23		−0.38	
	C.I.	(−0.64, 0.18)		(−0.98, 0.22)	
No. people/helping network (range: 0–10)	Comparison	5.30	(165)	4.86	(67)
	NV Pregnancy	5.16	(90)	4.30	(36)
	NV Pregnancy/Infancy	5.01	(99)	4.83	(37)
	Comparison-NV Infancy	0.29		0.03	
	C.I.	(−0.37, 0.95)		(−0.94, 1.00)	
No. intimates/helping network (range: 0–10)	Comparison	1.75	(165)	2.01	(67)
	NV Pregnancy	2.08	(90)	1.78	(36)
	NV Pregnancy/Infancy	1.57	(99)	1.62	(37)
	Comparison-NV Infancy	0.18		0.39	
	C.I.	(−0.22, 0.58)		(−0.23, 1.01)	
No. kin/helping network (range: 0–10)	Comparison	3.42	(165)	2.44	(67)
	NV Pregnancy	3.09	(90)	1.98	(36)
	NV Pregnancy/Infancy	3.09	(99)	2.09	(37)
	Comparison-NV Infancy	0.33		0.35	
	C.I.	(−0.18, 0.84)		(−0.40, 1.10)	
Expected accompaniment to labor and delivery (range: 0–16)[c]	Comparison	9.66	(165)	9.23	(67)
	NV Pregnancy	9.36	(90)	9.31	(36)
	NV Pregnancy/Infancy	9.04	(99)	8.69	(37)
	Comparison-NV Infancy	0.62		0.54	
	C.I.	(−0.12, 1.36)		(−0.17, 1.25)	
Grandmother support[d] (range: −7.10 to 16.10)	Comparison	0.12	(165)	1.17	(67)
	NV Pregnancy	−1.07	(90)	−1.56	(36)
	NV Pregnancy/Infancy	−0.07	(99)	1.33	(37)
	Comparison-NV Infancy	0.19		−0.16	
	C.I.	(−1.26, 1.64)		(−2.64, 2.32)	
Husband/boyfriend support[e] (range: −13.20 to 12.40)	Comparison	0.85	(165)	−4.90	(67)
	NV Pregnancy	1.55	(90)	−2.53	(36)
	NV Pregnancy/Infancy	0.34	(99)	−0.49	(37)
	Comparison-NV Infancy	0.51		−4.41	
	C.I.	(−1.52, 2.54)		(−7.37, −1.45)	

(continued)

TABLE 2—*(continued)*

Dependent Variable	Treatment Group	Whole Sample			Poor, Unmarried Women		
			$\bar{X}$	N		$\bar{X}$	N
		Odds			Odds		
Proportion husband boyfriend employed	Comparison	2.23	0.69	(150)	0.67	0.40	(43)
	NV Pregnancy	2.70	0.37	(83)	1.04	0.51	(32)
	NV Pregnancy Infancy	2.45	0.71	(92)	1.22	0.55	(35)
	Odds Ratio	1.10			1.82		
	C.I.	(0.58, 2.08)			(0.77, 4.44)		

[a]Values less than 59 indicate Hollingshead classes IV and V.
[b]Scale measuring extent to which women felt in control over their life circumstances, using a short-form variant of Rotter's locus of control instrument.[47]
[c]Scale summarizing women's responses to 2 Likert scales measuring their confidence that the individual they identified as support person would accompany them to labor room and to delivery. Higher scores indicate greater confidence.
[d]Scale characterizing availability, contact, and anticipated help with pregnancy and childrearing from the women's own mothers. Higher scores indicate greater support.
[e]Scale characterizing availability, contact, and anticipated help with pregnancy and childrearing from the women's husbands or boyfriends. Higher scores indicate greater support.

Note: Odds Ratio = $\frac{\text{NV Pregnancy Infancy}}{\text{Comparison}}$

cent reduction); this effect, however, did not extend into the two-year period following the end of the intervention at the 24th month postpartum.

Subsequent Pregnancies

During the first 22 months after delivery, nurse-visited, poor, unmarried women had one-third as many subsequent pregnancies as poor unmarried women in the comparison group (Table 5). Between birth and 46 months postpartum, this reduction in number of subsequent pregnancies was present for the sample as a whole (for which there was a 23 per cent reduction), but as before it was concentrated among the poor, unmarried women (for whom there was a 43 per cent reduction). These differences in number of subsequent pregnancies were reflected in (nonsignificant) reductions in the number of births and spontaneous abortions, and in an interval between the birth of the first and second child that was an average of 12 months longer.

Discussion

During the first four years after delivery of their first child, nurse-visited women who had not completed their high school education at the time they registered in the study returned to school more rapidly than their comparison-group counterparts; and nurse-visited women who were poor and unmarried were employed 82 per cent more of the time, had 43 per cent fewer subsequent pregnancies, and delayed the birth of their second child an average of 12 months longer. The effects of the program on employment were conditioned by the age of the mothers and the period during which assessments of employment were carried out. During the first two years postpartum, the nurse-visited, poor, unmarried older women worked more than their counterparts in the comparison group, but the nurse-visited, poor, unmarried teenagers did not do so. The nurse-visited, poor, unmarried teenagers were more concerned about eventually finding work, however, and during the two-year period after the program ended, they began to work more than poor, unmarried teens in the comparison group.

With a program designed to respond flexibly to the individual needs of socially disadvantaged families, it makes sense that the positive effects of the program would depend on the families' circumstances. That the impact of the program on educational achievement would be greater among women who had not graduated from high school is consistent with the program plan, in that the nurses emphasized the importance of the young women at least completing their high school education. That the effect of the program was only temporary suggests that many young women sought employment as soon as they were old enough to do so, irrespective of what impact this might have on their long-term educational achievement. While many of the nurse-visited women apparently shifted their focus to earning a living, their counterparts in the comparison group caught up with them educationally.

Unmarried, poor women are at increased risk for welfare dependence once they bear a child,[22] so it is with this particular group that one would expect the program to have the greatest opportunity for improvement in employment status. It makes sense that the nurses' success in helping unmarried, poor women find employment would be conditioned by the women's age, because employment opportunities for adolescents were limited and because a substantial number were still trying to complete their high school education. The increase in number of months that the nurse-visited poor, unmarried teenagers worked during the two-year period after the program ended may be explained by their increasing age and their having returned to school more rapidly after delivery. It is worth noting (Table 4) that while the nurse-visited poor, unmarried teens worked more than their counterparts in the comparison groups, both groups worked more during the two-year follow-up period than they had during the first two years after delivery. The difference in employment probably was influenced by the reduction in number of subsequent pregnancies, although it also may be argued that the nurse-visited women's educational and occupational aspirations and achievements contributed to the prevention of repeated pregnancy.[5,48]

The marked reduction in subsequent pregnancy among poor, unmarried women is consistent with the nurses' efforts to prevent *unintended* pregnancy. Although the extent to which pregnancy is unintended among unwed women and adolescents is unclear,[49] it is more likely to be unintended in these groups than among older, married women.[50,51] From the available data, we are unable to determine the extent to which this reduction in subsequent pregnancy reflects in-

TABLE 3—Estimates of Treatment Effects and 95% Confidence Intervals for Educational Achievement at 6, 10, 22, and 46 Months Postpartum (Adjusted for Covariates Husband/Boy Friend Support, Maternal Sense of Control, Maternal Age, and Social Class)

		Women < 12 Years Education at Registration			Unmarried Women < 12 Years Education at Registration		
Dependent Variables	Treatment Group	Odds	$\bar{X}$	N	Odds	$\bar{X}$	N
Proportion enrolled or graduated (6th month postpartum)[a]	Comparison	0.37	0.27	(52)	0.54	0.35	(42)
	NV-Pregnancy	0.35	0.26	(20)	0.22	0.18	(13)
	NV-Pregnancy/Infancy	1.44	0.59	(35)	2.23	0.69	(28)
	Odds Ratio	3.89			4.12		
	C.I.	(1.04, 15.48)			(1.20, 15.18)		
Proportion enrolled or graduated (10th month postpartum)[a]	Comparison	0.56	0.36	(51)	0.45	0.31	(41)
	NV-Pregnancy	0.56	0.36	(21)	0.47	0.32	(12)
	NV-Pregnancy/Infancy	1.08	0.52	(34)	1.50	0.60	(27)
	Odds Ratio	1.93			3.33		
	C.I.	(0.63, 6.25)			(0.96, 11.70)		
Proportion enrolled or graduated (22nd month postpartum)[a]	Comparison	0.61	0.38	(49)	0.67	0.40	(38)
	NV-Pregnancy	1.63	0.62	(20)	1.78	0.64	(12)
	NV-Pregnancy/Infancy	1.08	0.52	(32)	0.85	0.46	(26)
	Odds Ratio	1.77			1.27		
	C.I.	(0.49, 6.11)			(0.38, 4.06)		
Number of years education completed (46 month postpartum)	Comparison	—	11.13	(54)	—	11.46	(42)
	NV-Pregnancy	—	11.54	(21)	—	11.91	(12)
	NV-Pregnancy/Infancy	—	11.35	(41)	—	11.53	(33)
	Comparison-NV Infancy		−0.22			−0.07	
	C.I.		(−0.83, 0.39)			(−0.62, 0.48)	

[a]Dichotomous variable constructed to distinguish those women who at the time of interview had either graduated from high school, remained in high school, or enrolled in some other form of vocational training from those who had dropped out of high school before graduating.

Note: Odds Ratio = $\frac{\text{NV-Pregnancy/Infancy}}{\text{Comparison}}$

creased knowledge on the part of the nurse-visited women about how to prevent unintended pregnancy as opposed to their protecting their new prospects for educational and occupational achievement.

The presence of treatment effects concentrated on groups in greatest need in these areas is conceptually coherent, consistent with the program plan, and in keeping with at least one other intervention study addressing these issues.[28] The recent evaluation of Project Redirection, a national demonstration program aimed at helping young disadvantaged women attain economic self-sufficiency, found the strongest and most enduring effects to be concentrated among the most disadvantaged women.[28]

Several earlier studies also indicated that program effects on subsequent pregnancy, education, and employment disappeared within two to three years after the service ended.[23,28,29] In the current study, the corresponding pattern is mixed. For the sample as a whole, the impact of the program on subsequent pregnancy became stronger over time, while it began to diminish for the poor unmarried women. For those who had not yet graduated from high school, the influence of the program on educational achievement was short-lived; but for the poor, unmarried teenagers its impact on employment was not realized until after the program ended — once they became old enough to hold jobs.

Despite their coherence, these findings must be viewed with caution. Before the provision of services, the nurse-visited, unmarried women exceeded their comparison-group counterparts in sense of personal control, and the nurse-visited poor, unmarried teenagers exceeded their counterparts in boy friend support. Although we adjusted statistically for these preintervention differences, it is possible that there were other associated conditions that biased the sample in unknown ways. It should be emphasized, nevertheless, that this preintervention difference was attenuated because nurse-visited women with a greater sense of control discontinued the study more frequently than did their comparison-group counterparts. Moreover, no other preintervention differences were detected after a careful examination of the sample for a wide range of conditions that might bias relevant treatment contrasts.

Another reason for caution in interpretation is that the positive results of the current study were focused on specific at-risk subgroups. It may be argued that such findings have a greater chance of being sampling artifacts than do findings present for the sample as a whole. We do not subscribe to this interpretation, however, for two reasons. First, the effects detected for particular high-risk groups emerged from a series of planned comparisons incorporated in a set of statistical models that included factors (risk characteristics) identified at the stage of research design and used for sample recruitment (see Appendix). The second reason is the coherence of the findings.

There are a number of features of the present study that

TABLE 4—Estimates of Treatment Differences and 95% Confidence Intervals for Employment, Child Care, and Public Assistance Outcomes (Adjusted for Husband/Boy Friend Support and Maternal Sense of Control)

Dependent Variable	Treatment Group	Whole Sample		Poor, Unmarried Teenagers (<19 years at registration)		Poor, Unmarried Older Women (≥19 years at registration)	
		$\bar{X}$	N	$\bar{X}$	N	$\bar{X}$	N
No. months employed (0–22 months postpartum)	Comparison	5.89	(122)	2.17	(26)	3.61	(25)
	NV Pregnancy	7.57	(63)	1.56	(10)	5.41	(12)
	NV Pregnancy/Infancy	6.15	(72)	2.06	(15)	9.27	(10)
	Comparison-NV Infancy	−0.26		0.11		−5.66	
	C.I.	(−2.22, 1.70)		(−4.67, 4.89)		(−11.12, −0.20)	
No. months employed (0–46 months postpartum)	Comparison	12.65	(127)	10.03	(28)	7.14	(26)
	NV Pregnancy	17.01	(69)	5.03	(12)	15.08	(13)
	NV Pregnancy/Infancy	15.18	(81)	14.90	(16)	16.41	(12)
	Comparison-NV Infancy	−2.53		−4.87		−9.26	
	C.I.	(−7.29, 2.23)		(−13.89, 4.15)		(−19.45, 0.91)	
	Comparison-NV Infancy				−7.07		
	C.I.				(0.25, 13.89)		
Concern about finding work[a] (10th month postpartum)	Comparison	2.17	(128)	2.48	(27)	2.64	(25)
	NV Pregnancy	2.43	(67)	1.98	(11)	2.39	(12)
	NV Pregnancy/Infancy	2.38	(74)	3.44	(16)	2.26	(11)
	Comparison-NV Infancy	−0.21		−0.96		0.38	
	C.I.	(−0.58, 0.16)		(−1.69, −0.23)		(−0.45, 1.21)	
Concern about finding work (22nd month postpartum)	Comparison	1.92	(121)	1.90	(25)	2.65	(26)
	NV Pregnancy	2.24	(62)	1.97	(10)	2.51	(11)
	NV Pregnancy/Infancy	1.97	(71)	2.67	(15)	1.77	(10)
	Comparison-NV Infancy	−0.05		−0.77		0.88	
	C.I.	(−0.43, 0.33)		(−1.51, −0.03)		(0.05, 1.71)	
Help with child care[b] (10th month postparum)	Comparison	2.64	(130)	2.96	(28)	2.36	(26)
	NV Pregnancy	2.77	(67)	2.46	(11)	2.57	(12)
	NV Pregnancy/Infancy	2.76	(74)	2.77	(16)	3.17	(11)
	Comparison-NV Infancy	−0.12		0.19		−0.81	
	C.I.	(−0.42, 0.18)		(−0.39, 0.77)		(−1.47, −0.15)	
Help with child care (22nd month postpartum)	Comparison	2.76	(125)	2.68	(26)	2.52	(26)
	NV Pregnancy	2.94	(64)	2.83	(11)	2.70	(12)
	NV Pregnancy/Infancy	2.83	(72)	2.80	(15)	3.05	(10)
	Comparison-NV Infancy	−0.07		−0.12		−0.53	
	C.I.	(−0.40, 0.26)		(−0.77, 0.53)		(−1.26, 0.20)	
Help with child care (46th month postpartum)	Comparison	3.00	(107)	3.02	(23)	2.71	(17)
	NV Pregnancy	2.97	(60)	2.64	(9)	3.25	(10)
	NV-Pregnancy/Infancy	3.18	(76)	2.95	(13)	3.22	(12)
	Comparison-NV Infancy	−0.18		0.07		−0.51	
	C.I.	(−0.48, 0.12)		(−0.54, 0.68)		(−1.17, 0.15)	
No. days on public assistance (0–22 months postpartum)	Comparison	244	(122)	352	(52)	396	(26)
	NV Pregnancy	268	(63)	471	(22)	265	(12)
	NV Pregnancy/Infancy	236	(72)	419	(25)	239	(10)
	Comparison-NV Infancy	8		−67		157	
	C.I.	(−76, 92)		(−223, 89)		(−23, 337)	
No. days on public assistance (0–48 months postpartum)	Comparison	530	(127)	815	(54)	851	(26)
	NV Pregnancy	510	(69)	1020	(25)	443	(13)
	NV Pregnancy/Infancy	437	(81)	899	(28)	684	(12)
	Comparison-NV Infancy	93		−84		167	
	C.I.	(−78, 264)		(403, 235)		(−222, 556)	

[a]Likert scale indicating frequency of worrying about finding work (range: 1 to 4).
[b]Scale based on average frequency of help with specific child care tasks (e.g., diaper changing, dressing, toilet training, feeding, playing) provided by other family members and friends (range: 1 to 4.86).

TABLE 5—Estimate of Treatment Differences and 95% Confidence Intervals for Subsequent-Pregnancy Outcomes (Adjusted for Husband/Boy Friend Support and Maternal Sense of Control)

		Whole Sample			Poor, Unmarried Women		
Dependent Variable	Treatment Group	Log Incidence	$\bar{X}$	N	Log Incidence	$\bar{X}$	N
No. pregnancies (0–22 months)	Comparison	−0.81	0.44	(124)	−0.67	0.51	(52)
	NV-Pregnancy	−1.17	0.31	(64)	−0.76	0.46	(23)
	NV-Pregnancy Infancy	−0.97	0.38	(72)	−1.77	0.17	(25)
	Comparison-NV Infancy	0.16			1.10		
	C.I.	(−0.33, 0.65)			(0.15, 2.05)		
No. pregnancies (0–46 months)	Comparison	0.06	1.06	(126)	0.0	1.02	(54)
	NV-Pregnancy	−0.11	0.90	(68)	0.03	1.03	(24)
	NV-Pregnancy Infancy	−0.20	0.82	(79)	−0.54	0.58	(28)
	Comparison-NV Infancy	0.26			0.56		
	C.I.	(−0.04, 0.56)			(0.03, 1.09)		
No. spontaneous abortions (0–46 months)	Comparison	−2.08	0.13	(126)	−2.21	0.11	(54)
	NV-Pregnancy	−2.11	0.12	(68)	−1.87	0.15	(24)
	NV-Pregnancy Infancy	−2.30	0.10	(79)	−3.44	0.03	(28)
	Comparison-NV Infancy	0.22			1.23		
	C.I.	(−0.63, 1.07)			(−0.93, 3.39)		
No. therapeutic abortions (0–46 months)	Comparison	−3.18	0.04	(126)	−2.08	0.12	(54)
	NV-Pregnancy	−2.95	0.05	(68)	−1.80	0.17	(24)
	NV-Pregnancy Infancy	−3.02	0.05	(79)	−2.31	0.10	(28)
	Comparison-NV Infancy	−0.16			0.23		
	C.I.	(−1.23, 0.91)			(−1.15, 1.61)		
No. births including first born (0–46 mos)	Comparison	0.62	1.85	(133)	0.58	1.78	(58)
	NV-Pregnancy	0.57	1.77	(70)	0.55	1.74	(25)
	NV-Pregnancy Infancy	0.50	1.65	(83)	0.32	1.37	(31)
	Comparison-NV Infancy	0.12			0.26		
	C.I.	(−0.10, 0.34)			(−0.10, 0.62)		
No. months between first and second child[a] (0–46 mos)	Comparison	—	37.11	(133)	—	37.28	(58)
	NV-Pregnancy	—	41.58	(70)	—	42.40	(25)
	NV-Pregnancy Infancy	—	40.76	(83)	—	49.33	(31)
	Comparison-NV Infancy		3.65			12.05	
	C.I.		(−1.34, 8.64)			(4.48, 19.62)	

[a]Variable constructed to estimate interval between first and second birth. Those women who had not given birth to a second child and were not pregnant at the 46th-month interview were assigned interval values as if they gave birth 9 months after the date of the interview. Those women who were pregnant with their second child at the 46th-month interview were assigned interval values on the basis of their expected date of delivery.

limit its generalizability to other programs and communities: First, the program was carried out under favorable circumstances. The nurses were hired and trained exclusively for this experimental program, and each nurse carried a manageable caseload. Second, the community in which the research was carried out is not representative of hard-core inner cities or extremely isolated rural communities. Third, although the women and children enrolled in this study do represent a substantial portion of the women and children in the United States, there are many to whom these results cannot be applied (e.g., Blacks, those who register for prenatal care after 30 weeks of gestation). Consequently, the present study should be replicated in other settings, with other populations, and under different administrative arrangements before its findings are used as a basis for major policy or program initiatives in this area.

Although the value of this kind of comprehensive program may be questioned on the basis of its impact on maternal life-course development alone, the potential cost-savings due to increased maternal employment and reductions in unintended pregnancy are substantial. Moreover, when considered from the standpoint of its simultaneous improvement of other aspects of maternal and child functioning, [37,38] it becomes a public health strategy worthy of careful consideration. Our results suggest that home-visiting nurses, working closely with parents and existing health and human service providers, may be able to produce demonstrable improvements in many aspects of maternal and child functioning that heretofore have been the province of several separate services. With augmented training, reasonable caseloads, and focused efforts, community health nurses could play a decisive role in reducing many of the risks and unfavorable outcomes associated with childbearing among poor, unmarried women.

APPENDIX

Treatment Assignment

Families enrolled in the program were stratified by marital status, race, and geographic region (based on census tract boundaries). These families were assigned at random to one of the four treatment groups. At the end of the intake interview the women drew their treatment assignments from a deck of cards. Separate decks were used for groups cross-classified by the women's race, marital status at intake, and for Whites, the geographic region in which they resided. To ensure reasonably balanced subclasses, the decks were reconstituted periodically to overrepresent those treatments with smaller numbers of subjects, a procedure similar to Efron's biased-coin designs.[52] (Women in Treatments 3 and 4 subsequently were assigned on a rotating basis, within their stratification blocks, to one of five home visitors.)

There were two deviations from the random assignment of women to treatments.

- First, in six cases, women who enrolled were living in the same household as other women already participating in the study. To avoid potential horizontal diffusion of the treatment in the case of different assignments within households, the six new enrollees were assigned to the same treatment as their housemates.
- Second, during the last six months of the 30-month enrollment period, the number of cards representing Treatment 4 was increased in each of the decks to enlarge the size of that group and to enhance the statistical power of the design to compare the effectiveness of the infancy home-visiting program with that of Treatments 1 and 2 as a means of improving postpartum outcomes. Analysis of selected outcomes confirmed that this slight confounding of treatments with time did not affect the pattern of treatment effects reported here.

The Basic Model and Methods of Statistical Analysis

Dependent variables for which a normal distribution was assumed were analyzed in the general linear model; dichotomous outcomes (such as educational status at 6, 10, and 22 months) in the logistic-linear model (assuming a binomial error distribution); and low-incidence outcomes in the form of counts (such as number of subsequent pregnancies) in the log-linear model (assuming a Poisson distribution).

A core model was derived consisting of a $3 \times 2 \times 2$ factorial structure: treatments (1 and 2 vs 3 vs 4) × maternal marital status (unmarried vs married) × social class (Hollingshead classes IV and V vs I, II, and III), with all interactions among these factors; two covariates measured at registration (maternal sense of control and reported husband/boy friend support) were included as well. Maternal age was also included in each analysis. For variables with assumed normal distributions, age was specified as a classification factor with two levels (≤ 18 years vs > 18), as were its interactions with the three other factors. In the generalized cases (binomial and Poisson), estimation was more stable in models with fewer subclasses. Age, in these cases, was specified as a covariate.

The analyses of dichotomous dependent variables measuring whether women had either graduated from high school or were currently in some kind of educational program were carried out in the sample limited to those women who had not yet graduated from high school at the time of registration in the study. The model specified treatments, maternal marital status, and their interactions as classification effects, and age and SES as additional covariates.

A thorough investigation in all models was carried out for each covariate to examine its interactions with categorical variables. Interactions between a covariate and one or more categorical variables (nonhomogeneous regressions) indicate that certain tests of means depend on the covariate, in that a different test exists for each covariate value. This investigation was undertaken both to ensure correct interpretations of mean differences and to examine any substantive findings resulting from the interactions between continuous and categorical effects.[53]

Estimates and tests were adjusted for all covariates, classification factors, and interactions. Analysis was by our own computer programs; iterative weighted least squares was used in the generalized binomial logistic-linear and Poisson log-linear models. The means presented correspond directly to the tests: they are equally weighted averages of the smallest-subclass means, adjusted for the covariates. In the generalized case the analysis is carried out and estimates obtained in terms of the linearized form of the model — the logits (or log of the odds) in the logistic models, and the logs of the incidence in the log-linear models. The tables show the estimates in this form for the log-linear case; in the logistic case the estimates are converted to odds for individual groups and to odds ratios for the treatment contrasts. Confidence intervals were derived from the variance of the estimates and are placed on mean differences, the odds ratios, and on differences of logs of the incidence. All statements of treatment difference in the text are based on 95 per cent confidence intervals; the term "trends" refers to 90 per cent confidence intervals. Also shown in the tables in the column labeled $\bar{X}$ is the transformation of these estimates to probabilities (for the logistic case) and to incidence (for log-linear cases).

We have reported the treatment main effects and the effects for poor unmarried women in those analyses that included treatments, social class, and marital status as the classification factors. The treatment effects are shown separately for the younger and older poor unmarried women in those analyses in which maternal age was added as a classification factor.

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HRSA Funds $4 Million Home Health Program

More than $4 million was recently awarded by the Health Resources and Services Administration to five states under an innovative program designed to help low-income persons avoid unnecessary hospitalization by receiving health care in the home.

Grants were awarded to the states of Hawaii, Mississippi, North Carolina, South Carolina, and Utah. The states will implement the program through a variety of community-based organizations, and must contribute $1 to the program for every $3 in federal funds.

The program is intended to demonstrate how a multidisciplinary team of health and social service providers can effectively prevent unneeded hospitalization. Program funds may be used to pay for skilled medical and related health services, including those of the physician.

For additional information, contact Mr. Sis at HRSA, US Public Health Service, Department of Health and Human Services, 5600 Fishers Lane, Room 14-43, Rockville, MD 20857. Tel: (301) 443-3377.

DTP Immunization Status and Tetanus Antitoxin Titers of Mexican American Children Ages Six Months through Eleven Years

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Abstract: Data from the Mexican American portion of the Hispanic Health and Nutrition Examination Survey (HHANES), conducted in 1982–83, were analyzed for the number of diphtheria, tetanus, and pertussis (DTP) immunizations reported for Mexican American children 6 months–11 years of age and for levels of tetanus antitoxin titers in Mexican American children 4–11 years of age. In Mexican American children 6 months–11 years, 98.2 per cent had one or more DTP immunizations reported (95 per cent CI: 97.5, 98.9%); 85.1 per cent had three or more DTP immunizations reported (95 per cent CI: 83.2, 87.0%). The reported immunization coverage in Mexican American children was corroborated by the tetanus antitoxin titers which were above the minimum protective level (≥0.01 IU/ml) in 99.6 per cent of the 4–11 year olds. Using the American Academy of Pediatrics' (AAP) recommendations for the number of DTP immunizations, 61.1 per cent of the children 6 months–11 years of age had the age-appropriate number of immunizations (95 per cent CI: 58.5, 63.7%). AAP immunization completion rates were higher for children who: had a source of care reported (62.1 vs 44.3%; 95% CI of the difference: 7.1, 28.5); had insurance coverage (63.5 vs 56.1%; 95% CI of the difference: 2, 12.8); lived in a standard metropolitan statistical area (SMSA)-not central city as compared to SMSA-central city or not SMSA (66.6 vs 57.1%; 95% CI of the difference: 4.3, 14.7); and had 12 or more completed years of education for the head of the household (65.4 vs 58.3%; 95% CI of the difference: 1.8, 12.4). These four predictors were used to identify subgroups within the Mexican American population at high risk for incomplete immunization status. (*Am J Public Health* 1988; 78:1446–1450.)

Introduction

The health status and use of health care services by the Hispanic population in the United States have been questioned, because of the language barrier faced by many Hispanics and their lower income and education levels, relative to the general population.[1,2] To help clarify these issues, the National Center for Health Statistics (NCHS) conducted the Hispanic Health and Nutrition Examination Survey (HHANES) to assess some aspects of the health and nutritional status of the Hispanic population in selected areas of the United States.

In the HHANES, one surrogate measure of health care use was reported diphtheria, tetanus, and pertussis immunizations (DTP) and tetanus serology. DTP shots are given in connection with routine preventive care and thus reflect use of health care services. DTP immunizations, along with other childhood immunizations, have been identified as a 1990 Health Objective for the Nation.[3] Data from HHANES provide the first information on the attainment of this objective by Mexican-American children.

This paper focuses on the number of DTP shots reported in the Mexican American portion of the Hispanic Health and Nutrition Examination Survey (HHANES) for children 6 months–11 years of age and tetanus titer levels among children 4–11 years of age.

Methods

HHANES Survey Design

The HHANES was conducted among three Hispanic subgroups from July 1982 through December 1984 in three target areas of the United States: Mexican Americans in five southwestern states, Puerto Ricans in the New York City area, and Cuban Americans in Dade County (Miami), Florida.[4] The sample design of HHANES was a stratified, multistage, probability sample of persons 6 months through 74 years of age. Only data from the Mexican American portion of the survey are presented in this report.

The Mexican American portion of HHANES was conducted among civilian, noninstitutionalized persons of Mexican origin in parts of five southwestern states (Arizona, California, Colorado, New Mexico, Texas). The sample of 14 primary sampling units (PSUs), counties or small groups of counties, included 9,894 sample persons of whom 7,462 (75 per cent) were interviewed and examined. The weighted estimates presented in this report are representative of Mexican Americans in these five southwestern states. The weights for each survey participants included adjustments for nonresponse and noncoverage of geographic areas.

There were two levels of noncoverage of the Hispanic population in the southwest HHANES: 1) noncoverage of the Mexican American population residing in excluded counties in the Southwest with a small number and/or per cent of Mexican American residents (about 3 per cent); and 2) noncoverage of the Mexican American population residing in excluded census block groups or enumeration districts with small numbers of eligible Hispanics within the sample PSUs (about 10 per cent). Given the above two levels of noncoverage, the net coverage rate of the 1980 Mexican origin population in the Southwest was approximately 87 per cent.[4]

The sample weights were also post-stratified to Mexican American population totals for the Southwest as estimated by the US Bureau of the Census at the midpoint of the Mexican American portion of HHANES, i.e., March 1, 1983. Details of the HHANES sample design, selection process, operational plan, quality control procedures, and questionnaires have been previously documented.[4]

Data Collection

The HHANES included both a household interview and a three-hour medical examination conducted in a specially designed mobile examination center (MEC). During the household interview, information on children less than 12 years of age was obtained from the parents or other respon-

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TABLE 1—Percentage of Mexican American Children 6 Months–11 Years of Age with Reported Diphtheria, Tetanus, and Pertussis Immunizations

Age (yrs)	Sample Size	Population Estimate (in 000's)	One or more DTP Immunizations	Three or more DTP Immunizations	AAP Recommendations for DTP Immunizations[a]
0.5–11	2,562	2,221	98.2 (97.5,98.9)	85.1 (83.2,87.0)	61.1 (58.5,63.7)
0.5–1	364	329	95.3 (92.3,98.3)	65.8 (59.0,72.6)	65.8 (59.0,72.6)
2–6	1,152	997	98.1 (97.0,99.2)	85.4 (82.6,88.2)	61.3 (57.5,65.1)
7–11	1,046	895	99.5 (98.9,100*)	91.7 (89.4,94.0)	59.1 (55.0,63.1)

a) American Academy of Pediatrics recommendations:
0.5–1 year—at least 3 DTP immmunizations;
2–6 years—at least 4 DTP immunizations;
7–11 years—at least 5 DTP immunizations;
( ) = 95% Confidence Interval
* = Rounding.
SOURCE: NCHS, HHANES, 1982–83

sible adult household member. Upon completion of the household interview, participants were scheduled for an examination that included tetanus antitoxin titers for children 4 through 11 years of age.

Study Population

A total of 3,000 children 6 months–11 years of age were selected for the Mexican American portion of HHANES; 85.4 per cent completed the immunization section of the household interview; 78.5 per cent participated in the examination.

In the 4–11 year age group, 2,027 children were selected for the survey: 84.7 per cent completed the immunization section of the interview; 78.2 per cent participated in the examination. Tetanus titers were measured in 70.6 per cent of the examined children 4–11 years of age. Failure to obtain an adequate blood specimen was the primary reason for not performing the tetanus titer assay. The children having tetanus antitoxin titers measured were slightly older than the interview and examination sample (mean age 7.9 versus 7.5 years); they did not differ, however, on any of the other variables used in this study.

Variables

Tetanus antitoxin titers were determined by the Immunology Laboratory of the Division of Bacterial Diseases, Center for Infectious Disease, Centers for Disease Control, using the hemagglutination assay method[5] as part of a collaborative project with the Division of Maternal and Child Health, Health Resources and Services Administration, and NCHS. A protective level of tetanus antitoxin titer was considered to be 0.01 IU/ml or greater.[6,7]

Age was reported as year of age at last birthday. Immunization histories were obtained by recall during the household interview; the protocol did not require verification of responses by the participant's shot record. A child was considered to be up-to-date on immunizations if he/she had received the number of DTP shots recommended by the American Academy of Pediatrics (AAP): 6 months–1 year, at least three DTP shots; 2–6 years, at least four DTP shots; 7–11 years, at least five DTP shots.[8]

Source of care referred to the reporting of a usual place where routine health care or treatment of illnesses was received.

The poverty-income ratio (PIR) was calculated by dividing the total household income by the poverty threshold for a household of that size. The poverty thresholds, published by the Census Bureau,[9,10] are based on the total income needed to maintain a household on a nutritionally adequate food plan. Of the 2,562 in the study population, 194 were missing data on the PIR and were excluded from the analyses of PIR.

Education was defined by the completed years of education for the head of the household. Of the 2,562 in the study population, 138 were missing data on education and were excluded from the analyses of education.

Residence was originally categorized in three ways: standard metropolitan statistical area (SMSA)-central city, SMSA-not central city, and not SMSA. Preliminary analyses found the rate of immunizations similar between SMSA-central city (57 per cent) and not SMSA (57.5 per cent). As a result, these two categories were collapsed in further analyses.

Statistical Analysis

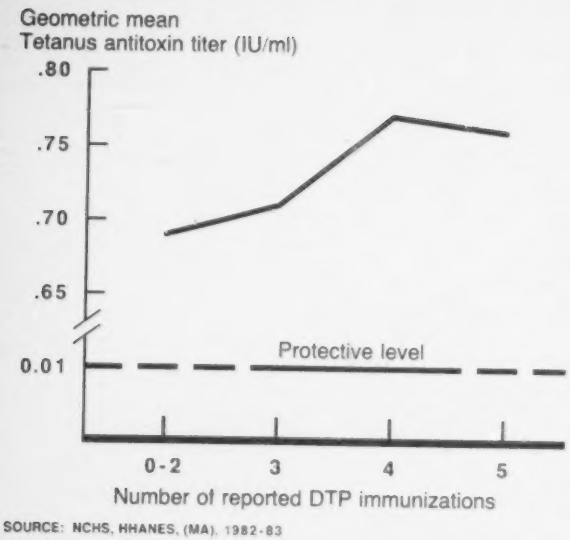
Sampling weights were used when calculating point estimates in order to account for individual selection probabilities and adjustments for non-response, non-coverage, and post-stratification which resulted from the complex survey design used in HHANES.

The complex survey design used in the HHANES tended to increase the variances which would have been obtained through simple random sampling.[11] The complex survey design was accounted for by using an average design effect to adjust the variances calculated under the assumption of simple random sampling. As described by Kovar,[12] the design effect is the ratio of the variance of a statistic from a complex sample to the variance of the same statistic from a simple random sample of the same size. The value of the design effect serves as an indication of the impact of the complex sample design on the variance. The average design effect for immunization status was 1.85. A t-test was used to compare point estimates and a linear test for trend was made using an orthogonal contrast.[13] All data analyses were done using programs accessible through SAS.[14]

Results

In Mexican American children 6 months–11 years of age, 98.2 per cent had one or more DTP immunizations reported (95% CI: 97.5, 98.9%); 85.1 per cent had three or more DTP immunizations reported (95% CI: 83.2, 87.0%) (Table 1).

This high level of reported immunizations is supported by the tetanus antitoxin titers of Mexican American children 4–11 years of age (Figure 1). Overall 99.6 per cent of Mexican American children had tetanus antitoxin titers above the



SOURCE: NCHS, HHANES, (MA), 1982-83

FIGURE 1—Tetanus Immunity Status of Mexican American Children four years through 11 years of age

minimal protective level of 0.01 IU/ml regardless of the number of DTP shots reported. A slight but unimportant increase in tetanus titers was noted with an increasing number of reported DTP shots. In the 4–11 year age group, 99.3 per cent had one or more DTP immunizations.

In the HHANES, using the AAP recommendations, as defined earlier, a more pessimistic picture emerges; 61.1 per cent of Mexican American children 6 months–11 years of age had up-to-date immunization coverage (95% CI: 58.5, 63.7%) (Table 1).

The association between various sample person characteristics and reported up-to-date immunization status according to AAP standards is shown in Table 2. Higher levels of up-to-date immunization coverage were found among children having a source of care reported, having health insurance coverage, living in a SMSA-not central city as compared with SMSA-central city or not SMSA, and with 12 or more completed years of education for the head of household. Higher levels of reported up-to-date immunization coverage were found in children living at or above the poverty level and those in families of four or fewer persons, however, these differences included 0 in the 95 per cent confidence interval of the difference. Little variation in up-to-date immunization coverage by sex was noted.

Four predictors of up-to-date immunization coverage were found in this study (source of care, health insurance, residence in a SMSA-not central city, and 12 or more completed years of education for the head of the household). These four variables were tested to see if they could be used to identify a segment of the population at high risk for incomplete immunization status (Table 3). When 0–1 predictors were present nearly 50 per cent of the children had incomplete immunization coverage. But when all four predictors of up-to-date immunization coverage were present, the percentage of children reporting incomplete immunization coverage dropped to 28.3 per cent.

Discussion

Virtually all 6 months–11 year old Mexican American children had at least one DTP immunization. The high level of immunization coverage found in HHANES may reflect the efforts directed at increasing immunization levels beginning with the 1977 US Initiative on Childhood Immunization.[15] As a result of this Initiative, all 50 states now have requirements for immunization at the time of entry to school.[16] All children in HHANES 6–11 years of age reported school attendance.

The tetanus antitoxin titers measured on 4–11 years old Mexican American children corroborated the high level of reported immunizations. The tetanus antitoxin titers reflect past tetanus immunization, since naturally acquired immunity to tetanus is virtually nonexistent in the United States, and the disease state itself does not confer immunity.[17]

Certain subgroups of Mexican American children, such as children of migrant workers, however, may not have been well covered in the survey. The rate of immunization may be lower for these subgroups than for those reported in the total Mexican American population.

The 1990 Objective for the Nation on Immunization Status sets as its goal at least 90 per cent of two-year-old children with their basic immunization series complete.[3] In Mexican American children 2 years of age, 76.4 per cent had received three or more DTP shots. Thus, in regard to DTP immunizations, Mexican American children in the Southwest are short of the 1990 Objective.

Data on childhood immunization levels were not collected in previous Health Examination Surveys. The only national data on immunization levels for a comparable time period are from the 1982 United States Immunization Survey (USIS), a supplement to the Current Population Survey.[18,19] These data, however, are not strictly comparable to the HHANES because of differences in survey methodology. Also, the USIS was a subsection of a survey dealing with labor force participation and not a health survey as HHANES was.

The data collected in the USIS was for two groups: the total sample and the record sample, a subgroup claiming to have based its reported immunization history on a shot record. The record sample was approximately 33 per cent of the total sample. The record sample was found to be more consistent with the state data.[20,21] The accuracy of the immunization levels reported in the USIS have been verified by comparing the results to state reports on immunization status of children entering school.

Comparison of the data from the USIS[22] with comparable age groups in the HHANES revealed somewhat confusing results. In the 1–4 year age group, 78.6 per cent of Mexican American children in HHANES had three or more DTP immunizations. In the 1982 USIS total sample, 70.6 per cent of White children and 48.4 per cent of children of Other Races had three or more DTP immunizations and in the 1982 USIS record sample 88.4 per cent of White children and 74.6 per cent of children of Other Races had three or more DTP immunizations. In the 5–9 year age group, 92.1 per cent of Mexican American children in HHANES had three or more DTP immunizations, while in the 1982 USIS total sample 79.3 per cent of White children and 63.4 per cent of children of Other Races had three or more DTP immunizations. In the 1982 USIS record sample, 95.1 per cent of White children and 87.0 per cent of children of Other Races had three or more immunizations.

Thus, in regard to DTP immunization status, the relative position of Mexican American children, based on HHANES data, in comparison to White children in the 1982 USIS changes depending on which sample is used. However, the

TABLE 2—Per Cent of Mexican American Children 6 Months–11 Years of Age with Reported Up-to-date Diphtheria, Tetanus, and Pertussis Immunizations[a] by Selected Characteristics

Characteristics	Size of Study Population	Population Estimate (in 000's)	Per Cent Immunized	95% Confidence Interval of the Difference
Total	2,562	2,221	61.1 (1.31)	
Reported Source of care				
Yes	2,404	2,090	62.1 (1.35)	
No	158	131	44.3 (5.28)	7.1, 28.5
Health Insurance				
Yes	1,700	1,495	63.5 (1.60)	
No	862	726	56.1 (2.27)	2.0, 12.8
Residence				
SMSA-not central city	1,048	926	66.6 (2.00)	
SMSA-central city or not SMSA	1,514	1,295	57.1 (1.71)	4.3, 14.7
Completed years of education for the head of the household[b]				
12+	1,014	919	65.4 (2.08)	
0–11	1,410	1,184	58.3 (1.75)	1.8, 12.4
Poverty Index Ratio (PIR)[b]				
At/Above PIR	1,490	1,324	62.5 (1.73)	
Below PIR	878	738	57.0 (2.23)	−0.03, 11
Number of persons in family				
2–4	1,077	966	63.4 (2.03)	
5–18	1,485	1,255	59.3 91.71)	−1.1, 9.3
Sex				
Male	1,279	1,117	60.2 (1.86)	
Female	1,283	1,104	62.0 (1.84)	−3.4, 6.9

a) American Academy of Pediatrics recommendation
b) Of the 2,562 in the study population, 194 were missing data on PIR and 138 were missing data on education. These individuals were excluded from the analyses for these respective variables.
( ) = Standard error
SOURCE: NCHS, HHANES, 1982–83

TABLE 3—Predictors[a] of Reported Up-to-date Immunization Coverage[b] in Mexican-American children 6 months through 11 years of age

Presence of Predictors of Up-to-Date Immunization Coverage	Size of study population[c]	Population Estimate (in 000's)	Per Cent with Incomplete Immunization Coverage
0–1 Predictors	447	379	49.4 (3.18)
2 Predictors	773	638	41.0 (2.35)
3 Predictors	839	740	35.8 (2.27)
4 Predictors	365	347	28.3 (3.36)*

a) Predictors of up-to-date immunization coverage: Source of care, health insurance, residence in an SMSA—not a central city, 12 or more completed years of education for the head of household.
b) American Academy of Pediatrics recommendations.
c) Of the 2,562 in the study population, 138 were missing data on education. These individuals were excluded from this analysis.
* = Test for trend $p<0.01$.
( ) = Standard error.
SOURCE: NCHS, HHANES, 1982–83.

relative positive of Mexican American children in comparison to children of Other Races does not vary between the two samples.

Examining the percentage of Mexican American children with reported up-to-date immunization coverage according to the AAP recommendations reveals a more pessimistic picture. Only 61.1 per cent of Mexican American children were considered up-to-date by these standards. However, with the addition of one more immunization, 79.8 per cent would be up-to-date and with the addition of two more immunizations, 93.0 per cent would be up-to-date. Thus, only two visits would be required to bring over 90 per cent of the children up-to-date on their immunizations.

The predictors of immunization coverage among Mexican American children in this study are similar to previously identified predictors: use or contact with the health care system,[23,25] socioeconomic status, e.g., income[26,27] and education[26–29]; residency in a certain area[26,27]; and family size.[26,27,29]

Identifying high-risk groups is the first step in the primary prevention of disease. In Table 3 the effectiveness of four predictors of immunization status are shown. These

predictors can be used by health care providers to identify Mexican American children at higher risk for incomplete immunization coverage.

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American Federation of Teachers to Participate in Immunization Awareness Program

The American Federation of Teachers (AFT) has joined with the national network of children's hospitals in efforts to arrest the decline in childhood vaccinations. As part of an immunization awareness program, the groups will disseminate information in inner-city areas that have low levels of immunization. Over the last decade, the immunization rates have declined among preschool children of immigrant, poor, and minority children, due primarily to the lack of knowledge in these populations regarding the value of vaccinations.

The program will provide information to parents so they understand how their children can be protected from the dangerous childhood diseases that can threaten unvaccinated children. Central to the effort is a new booklet, available in both Spanish and English, which was prepared through a grant from Lederle Laboratories. The booklet discusses the preventable diseases (polio, measles, mumps, rubella, pertussis, diphtheria, tetanus, and haemophilus type-B infections); tells parents where to get vaccines, how they are administered, possible side effects; and includes a checklist for keeping track of immunizations. The brochure also alerts parents to the availability of free vaccinations — encouraging families with financial hardship to contact the Children's Hospital National Medical Center in Washington, DC at (202) 879-4458 for more information.

Copies of the brochure can be obtained by sending a self-addressed, stamped envelope to AFT Order Department, 555 New Jersey Avenue, NW, Washington, DC 20001.

Drinking Water Source, Diarrheal Morbidity, and Child Growth in Villages with Both Traditional and Improved Water Supplies in Rural Lesotho, Southern Africa

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Abstract: This study examined the growth and morbidity rates of young children in relation to exclusive and non-exclusive use of improved water supplies in rural Lesotho, southern Africa. Data were collected for 247 children 60 months of age and under between July 1984 and February 1985 in 10 villages that had an improved water supply at least one year prior to investigation. Children whose families relied exclusively on the new water supply for their drinking and cooking needs grew 0.438 cm and 235 g more in six months than children whose families supplemented the new water supply with the use of contaminated traditional water for drinking and cooking. The difference in growth was greater among children over 12 months of age at the start of the evaluation than among infants. This may be explained partly by lower rates for *Giardia lamblia*, the most commonly identified pathogen in stools in older children. Among infants, similar rates of *Campylobacter*, the most commonly isolated pathogen among infants, may have prevented larger differences. Results suggest that improved drinking water supplies can benefit preschool children's health after infancy, but only if they are functioning and utilized exclusively for drinking and cooking purposes. (*Am J Public Health* 1988; 78:1451–1455.)

Introduction

The evidence for increased health benefits following improvements in the quality of drinking water is equivocal.[1] Young children in urban areas appear more likely to benefit from improved water supplies than young children in rural areas, perhaps because in urban areas drinking water constitutes a major source of exposure to pathogens, while in rural areas other sources of exposure may be more important.

In a previous study in rural Lesotho, southern Africa, improved water supplies were not associated with less diarrhea.[2] In a more recent health impact evaluation in Lesotho, children from villages that had access to an improved water supply had lower stool isolation rates for *Giardia lamblia* and *Escherichia coli*, but they did not have less overall diarrhea.[3] Growth among children from villages without an improved water supply was better than among children from villages with an improved water supply.[3]

Several factors[1,4] could account for the lack of more positive findings in these and other similar studies. One reason postulated, but never documented, is that the improved water supply was not the only water source used for drinking and cooking purposes. This study estimates the protective effect of exclusive versus non-exclusive use of improved water supplies on the prevalence of pathogens in stools, diarrhea rates, and growth of preschool children who resided in villages that had access to an improved water supply.

Methods

The study was conducted in Lesotho, which is a mountainous kingdom surrounded entirely by the Republic of South Africa. The country ranges in elevation from 1,500 to over 3,400 meters, with the majority of the population in the lowlands under 2,000 meters. Ninety per cent of the population (1.5 million) is rural despite increasing urbanization.

Ten villages were studied from July 1984 to February 1985, all of which received an improved water supply between 1967 and March 1983. They were representative of all the villages that received an improved water supply during this time in terms of size and geographic location. The villages were located in four districts that included 50 per cent of the rural population of Lesotho and were in the lowlands or foothills.

All villagers had access to a continually functioning tap or hand pump for at least one full year prior to data collection. The community taps or hand pumps, each of which served 100 people or less, were dispersed throughout the villages. Gravity fed springs, in which water passed through a silt box, supplied water to the taps. In the flatter areas, bore holes were dug when springs or other surface water was unavailable. Depth of the average bore hole was 50 meters, sometimes exceeding 100 meters. Over 90 per cent of the hand pumps installed are the Moyno pump.

Data used in these analyses were collected in three five-week periods: July–August 1984, the cold and dry season, October–November 1984, and January–February 1985, the warm and wet season. Of the 294 children, 1–60 months of age, enumerated during period one, 258 children were found during period 3 (87.3 per cent). Attrition was similar in both groups. Reasons for attrition were: 13 could not be accounted for, 10 were visiting out of the village, eight migrated to another village, parents of three children refused continuation in the study, and two died (one from each group). Finally, 11 children were dropped from the analyses because they never used an improved water supply, leaving a total of 247 children available for analysis.

During the first and third rounds, mothers were interviewed and children were measured for length and weight. The child's age in months at the start of the study was also taken from clinic cards and verified by use of a local events calendar.

Children were weighed and measured for length according to standard techniques,[5] using a portable CMS weighing

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scale (25 kg × 100 g) and measuring board (130 cm × 1 mm), respectively. Intraobserver (0.005 cm^2 and 0.001 kg^2) and interobserver (0.003 cm^2 and 0.001 kg^2) measurement error rates were small.[6]

In order to estimate growth differences for all 247 children over the six-month period final length or weight for all 247 children was regressed on the initial value,[7–11] and residuals from these regressions were then used as the growth outcome in subsequent analyses. Residual analysis accounts for all the effects that influence a child's body size at the start of the evaluation when examining the determinants of growth that occurred only during the time of the evaluation. Residuals are independent of the initial value and may be positive or negative. The difference in residual values between the exclusive and mixed users reflect differences in six-month growth among children. Results were similar when final body size was used as the outcome, while controlling for initial body size.

During round 2, stool samples were collected from 123 children, and they were examined for bacteria, parasites, and viruses according to standard methods reported elsewhere.[3] During all three rounds mothers were asked to report on the presence or absence of diarrhea in their children in the previous 24 hours. We used the common word "letsollo" in Sesotho, which means "running stomach."

Dummy variables related to use of water and socioeconomic conditions were constructed from the interview data. Children were classified according to whether the family relied exclusively on the improved water supply for their drinking and cooking needs (exclusive users) or supplemented the improved water supply with traditional water sources (mixed users). This information was obtained by interview during round 1. Children were also classified according to the family's per capita use of water drinking rounds 1 and 3. Changes in the quantity of per capita water use between rounds 1 and 3, increased versus decreased water use or no change, were also included in the analyses. Changes in per capita water use was a better predictor of growth than was per capita water use at either round 1 or round 3.[3]

Other potential confounding variables were included in the analyses. The dummy variables were village of residence, remittance from at least one person working in the Republic of South Africa (no, yes), latrines (no, yes), maternal literacy (no, yes), maternal marital status (no, yes), pregnancy (no, yes), frequency of maternal bathing (four or more times per week versus fewer than four), child's sex (male, female), child's age (1–12, 13–24, 25–36, 37–48, 49–60 months), and current breastfeeding (no, yes). Continuous variables were natural log of major household expenditures, number of pieces of agricultural equipment, number of major household possessions, household size, crowding (people/room), and parity. Some factors were kept as continuous variables because they explained more of the growth than if they were categorized.

Statistical analyses were done in SYSTAT, version 3.0,[12] a general linear program for least squares analysis of multivariate data. The child was the unit of analysis, and the variables listed above were included in the multivariate models. Ninety-five per cent confidence intervals were calculated. The effects on weight and length were adjusted for potential confounding by first including all factors above and then by including all factors except the effects due to changes in the amount of water used over seasons. Calculations of nutritional status of children compared to National Center for Health Statistics (NCHS) reference data were done using CASP, a software package from the Centers for Disease Control.[13] Length and weight gain, in separate analyses, were examined by comparing the odds ratio of being in the upper versus lower 25th quartile of growth for mixed and exclusive users.

Results

Of the 247 children, 125 were exclusive users and 122 were mixed users. Most potential confounders were distributed similarly between the two groups (Table 1). Three variables (remittance, maternal literacy, and crowding) were more prevalent among the exclusive users; they all were associated negatively with growth. Breastfeeding was also more common among exclusive users, but this was associated positively with growth only among infants. Frequent maternal bathing and increased water use were more common among mixed users; these factors were associated positively with growth. The percentage of exclusive users within each village was variable, ranging from 17.6 to 90.3.

Weight and length were similar between exclusive and mixed users at the start of the study (Table 2). Because initial body size and growth among children aged one, two, three, and four years was similar, only the combined effects of children 13–60 months of age at the start of the evaluation are reported. Mild malnutrition (low weight-for-age) in both groups was a result of stunting (low height-for-age), not wasting (low weight-for-height). In early infancy children followed the NCHS 50th percentile, but by the end of infancy they were already stunted and had lower weight-for-age

TABLE 1—Means of Potential Confounding Factors among Study Children by Use of Drinking Water

Factors	Exclusive Users (n = 125)	Mixed Users (n = 122)
Village of Residence	%	%
Ha Phoku (350)†	35.3	64.7
Phamistone (500)	17.6	82.4
Letsoare (300)	44.4	55.6
Qalabane (600)	24.0	76.0
Khabos (1500)	31.8	68.2
Matube (700)	82.1	17.9
Makhosi (300)	57.9	42.1
Lebone (700)	25.0	75.0
Mangnaneng (500)	87.5	12.5
Khatleng (350)	90.3	9.7
Household Characteristics		
Agricultural tools (mean of 5 items)	1.4	1.2
Crowding (people/room)	2.8	2.3
Household size	5.7	5.7
Ln household expenses (Maluti)	4.3	4.3
Possessions (mean of 12 items)	7.5	7.0
Remittances (yes)	86.4	79.5
Increase water use (yes)	16.0	22.1
Latrines (yes)	29.6	32.8
Maternal Characteristics		
Literacy (yes)	98.4	91.0
Pregnancy (yes)	16.0	16.4
Bathing (>4 times/week)	17.6	25.4
Married (yes)	86.4	92.6
Parity	3.4	3.5
Child Characteristics		
Age		
1–12 months	26.4	21.3
13–60 months	73.6	78.7
Breastfed (yes)	47.2	36.1
Male	48.8	46.7

†Village population size in parentheses

TABLE 2—Per Cent of NCHS Median Reference among children 1–60 Months of Age at Start of Study by Use of Drinking Water

Children's Age in Months	Exclusive Users		Mixed Users	
	Sample Size	NCHS ± SD	Sample Size	NCHS ± SD
		Weight for Age		
1–12	33	95.6 ± 13.5	26	95.7 ± 12.2
13–60	92	89.6 ± 9.4	96	88.3 ± 11.8
		Length for Age		
1–12	33	94.4 ± 5.1	26	94.9 ± 3.7
13–60	92	93.7 ± 3.7	96	93.7 ± 4.6
		Weight for Height		
1–12	33	110.4 ± 11.2	26	108.7 ± 9.4
13–60	92	100.4 ± 7.9	96	98.8 ± 8.3

values. Children within 13–60 months of age maintained their growth at about 88–89 per cent of the reference weight-for-age and 93–94 per cent of the reference length-for-age; wasting was not apparent.

Absolute unadjusted growth over the six-month period was similar between groups among the infants, but was greater among exclusive users compared with mixed users among the older children (Table 3). The difference in six-month growth between groups based on absolute growth or residuals was also similar. Only differences in six-month growth residuals between groups are presented below.

The average unadjusted difference in weight gain between exclusive and mixed users was 118 g (Table 4). Among infants, the unadjusted difference was inconsequential, but among children 13–60 months of age, the difference was 156 g. On average exclusive users also grew 0.236 cm more in length than mixed users (Table 5), 0.244 cm more among older children.

After adjustment for confounding, the weight gain of children was better among exclusive users than among mixed users whether or not the effects of changes in water quantity were controlled (Table 4). Exclusive users gained 235 g more than mixed users when changes in water quantity were controlled and 226 g when it was not controlled. The difference in length gain between exclusive and mixed users of the new water supply was 0.438 cm when water quantity was controlled and 0.404 cm when it was not controlled (Table 5). Among older children the effects were somewhat

TABLE 3—Six-month Growth among Children by Age at Start of Study and Use of Drinking Water

Children's Age in Months	Exclusive Users		Mixed Users	
	n	Growth ± SD	n	Growth ± SD
		Weight (g)		
1–12	33	1.036 ± 0.691	26	1.035 ± 0.355
13–60	92	0.817 ± 0.456	96	0.668 ± 0.514
		Length (cm)		
1–12	33	6.464 ± 2.093	26	6.338 ± 1.413
13–60	92	4.239 ± 1.149	96	3.949 ± 1.215

TABLE 4—Differences in Six-Month Weight Gain between Exclusive and Mixed Users of Improved Drinking Water Supplies among 247 Study Children.

Children's Age in months	Weight Differences (g)		
		Adjusted for Confounding†	
		Water Quantity	
	Unadjusted (95% CI)	Controlled (95% CI)	Not Controlled (95% CI)
1–12	1 (−244,246)	1 (−258,260)	24 (−231,279)
13–60	156 (20,292)	293 (114,472)	276 (96,456)
Total	118 (−2,238)	235 (77,394)	226 (66,386)

†All variables listed in the methods section were controlled.

larger for weight and length gain when changes in water quantity were controlled than when it was not controlled; among infants the effects were smaller. When the odds ratio of being in the upper versus lower quartile was examined, mixed users were twice as likely to be in the lowest quartile than exclusive users for both length and weight gain.

In a subsample of 123 children in this study, stool isolation rates were 33 per cent lower for *G. lamblia*, 53 per cent lower for enteropathogenic *E. coli*, and 49 per cent lower for coronavirus-like particles among exclusive users compared with mixed users (Table 6). These pathogens were more common among older than younger children. *Campylobacter* was only slightly less common among infants in the exclusive group compared with the mixed group and *Giardia* was rare in infancy. Rotavirus, *Salmonella* and *Shigella* were not detected. Diarrhea rates were reported to be generally higher among exclusive users than mixed users (Table 7), but differences between groups were small except among infants during round 1.

Discussion

In summary, exclusive use of improved water supplies for drinking and cooking was associated with better child growth, depending on the child's age. Source of drinking water made virtually no difference among infants; among one- to four-year old children, it did affect growth. By extrapolation, the expected improvement in growth over the first five years of life would be 4.4 cm and 2.3 kg.

Comparison groups were similar in that only children from villages with an improved water supply were included in the analyses, thus removing any biases due to systematic differences between villages with and without improved water supplies. Exclusive users still grew less well, however, than children from non-improved villages,[6] even though they grew better than mixed users. Improved and unimproved villages were similar in rurality, village size, presence of schools or clinics, age and sex distribution, and general standard of living. Still, selection bias in the comparison of villages with and without improved water supplies cannot be ruled out, even though virtually all villages applied for a water supply. For instance, water supplies were constructed deliberately in villages that had a recent outbreak of disease.

In the comparison of exclusive and mixed users within villages with improved water supplies, uncontrolled differences may have existed such as distance to the tap or number of people for each tap. Those households nearest to the tap

TABLE 5—Differences in Six-Month Length Gain between Exclusive and Mixed Users of Improved Drinking Water Supplies among 247 Study Children.

Children's Age in months	Length Differences (cm)		
		Adjusted for Confounding[1]	
		Water Quantity	
	Unadjusted (95% CI)	Controlled (95% CI)	Not Controlled (95% CI)
1–12	0.107 (−0.629, 0.843)	0.083 (−0.576, 0.742)	0.174 (−0.494, 0.842)
13–60	0.244 (−0.103, 0.591)	0.527 (0.072, 0.982)	0.462 (0.000, 0.924)
Total	0.236 (−0.087, 0.559)	0.438 (0.035, 0.841)	0.404 (−0.006, 0.814)

[1]All variables listed in the methods section were controlled.

TABLE 6—Stool Isolation Rates for Selected Pathogens among Children 1–60 Months of Age at Start of Study by Use of Drinking Water

Pathogen	Exclusive Users		Mixed Users	
	n	Number (%)	n	Number (%)
Campylobacter				
1–12 months	20	5 (25.0)	14	4 (28.6)
13–60 months	40	3 (7.5)	49	0 (0.0)
Escherichia coli enteroinvasive				
1–12 months	20	0 (0.0)	14	1 (7.1)
13–60 months	40	1 (2.5)	49	0 (0.0)
enteropathogenic				
1–12 months	20	1 (5.0)	14	2 (14.3)
13–60 months	40	3 (7.5)	49	7 (14.3)
enterotoxigenic				
1–12 months	20	3 (15.0)	14	0 (0.0)
13–60 months	40	3 (7.5)	49	3 (6.1)
Giardia lamblia				
1–12 months	20	1 (5.0)	14	1 (7.1)
13–60 months	40	8 (20.0)	49	13 (27.1)
Coronavirus-like particles				
1–12 months	20	3 (15.0)	14	3 (21.4)
13–60 months	40	10 (25.0)	49	24 (49.0)

or hand pump may have been different from the mixed users in resources or attitudes toward health, allowing them to be exclusive users. For example, these households may have had substantial influence in the village, thereby securing not only the power to determine where the new supplies were located but also the ability to mobilize limited resources that also fostered better child growth. Energy intakes between exclusive and mixed users also were not controlled and cannot be ruled out as a potential confounding effect.

The classification of exclusive versus mixed users was done by maternal reporting, not by observation. No incentive was apparent for mothers to report mixed use if they were in fact strictly exclusive users. On the other hand, mixed users may have reported exclusive use since engineers told villagers that the improved water supply should be used solely for drinking and cooking purposes and not for domestic hygiene. If such misclassification did exist, the differences reported above would be underestimated.

Improved water supplies were continually functioning and available for use during the time of the study and for at least one year prior to the study, with three exceptions:

- Prior to data collection one tap from one village had not been working for several weeks. It was reported and fixed prior to data collection.
- Water flow from improved supplies was limited occasionally either because of inadequate source water or because of volunteer restriction from the villagers themselves. In some instances, the taps were shut down for several hours in the middle of the day. Since most water in Lesotho was usually collected during morning and evening hours,[2] the midday closure of these taps probably did not explain differences in patterns of use.
- In another village, one tap was broken for several months and was not fixed. Ten children came from families that lived near this tap and did not use alternative improved water supplies elsewhere in the village. In yet another village, a child was not allowed to use the tap because the family did not contribute with either labor or to a general maintenance fund while receiving the improved water supply. These eleven children were excluded from the analyses because they may have been different from the other children. They gained less in weight but more in length than either the exclusive or mixed users.[6]

Possible reasons for the lack of more positive results among infants was the finding of *Campylobacter* almost solely among infants and the similar isolation rates among both groups. On the other hand, among older children higher isolation rates of *Giardia lamblia*, particularly among mixed users, may have accounted for part of the growth differences between groups. The association between coronavirus-like particles and diarrhea has not been adequately demonstrated, but the lower rate of coronavirus-like particles among exclusive users may have contributed to better growth.

Reported diarrhea rates were actually higher among

TABLE 7—Diarrhea Rates among Children 1–60 Months of Age at the Start of the Study by Use of Drinking Water

Children's Age in months	Exclusive Users		Mixed Users	
	n	Number (%)	n	Number (%)
Period 1				
1–12	33	4 (12.1)	26	0 (0.0)
13–60	92	8 (8.7)	96	4 (4.2)
Period 2				
1–12	29	4 (13.8)	19	3 (15.8)
13–60	81	5 (6.2)	74	3 (4.1)
Period 3				
1–12	33	8 (24.2)	25	4 (16.0)
13–60	87	5 (5.7)	90	3 (3.3)

exclusive users, but reporting was subject to recall bias, defined by the mother, and not verified by clinical examination or observation. The growth variables, on the other hand, were measured with minimal error and not subject to misreporting.

Among infants growth differences between exclusive and mixed users were greater when per capita water use was not included in the analyses and virtually disappeared when water use was controlled, suggesting that among infants, use of water for cleanliness is probably more important than is clean drinking water, probably because virtually all infants were breastfed. The reverse was true among older children. Water quality may become more important as the protective effect of breastfeeding is removed.[14]

Larger villages tended to have a lower percentage of exclusive users than smaller villages. If village size, number of people per tap, or distance to taps was large, continued use of contaminated supplies may have occurred. The installation of more taps closer to individual homes might perhaps avoid the use of contaminated supplies and foster the exclusive use of the improved water supplies for drinking and cooking purposes.

Previous reports in Lesotho[2,3] indicate that improved water supplies must be coupled with health education to encourage people to use only the improved water supplies for drinking and cooking purposes. These supplies must also be maintained to ensure continual functioning, a crucial requirement often missing in water supply projects. Future studies of health impacts from improved water supplies should measure factors affecting water use such as distance to tap, number of families per tap, and reliability of the operation of the new or improved water supplies.

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NEHA 53rd Annual Educational Conference: Call for Abstracts

The National Environmental Health Association (NEHA) has issued a call for abstracts for its 53rd annual educational conference to be held in Seattle, Washington June 24-29, 1989. The conference theme — The Nineties: A Decade for Change — will focus on the rapidly changing technologies and methods used to safeguard the public health from environmental threats. Special features of the conference will include:

- A computer software exchange featuring the latest innovations in software for the environmental health professional
- Two pre-conference seminars dealing with asbestos removal and the offering of a food certification examination
- A special all-day theme session dealing with public relations and the environmental health/protection professional

NEHA — a professional society composed of environmentalists from all areas of the field, including public and private sectors and academe — invites abstracts for papers to be presented at the June conference. Interested participants should submit a typed one-page abstract outlining the proposed presentation. The deadline for submissions is November 30, 1988.

For further information, contact Mel Monkelis, conference manager, NEHA, 720 S. Colorado Boulevard, Room 970 South Tower, Denver, CO 80222. Tel: (303) 756-9090.

Seating Position in Cars and Fatality Risk

LEONARD EVANS, DPHIL, AND MICHAEL C. FRICK, MA

Abstract: Fatality risk in passenger cars according to seating position (front versus rear; left versus center or right) was examined using Fatal Accident Reporting System (FARS) data for 1975 through 1985. Comparing the fatality risk of unrestrained occupants matched in sex and age (within three years) revealed effects attributable to seating position, and not to occupant characteristics correlated with use of different seats. Fatality risk to drivers was the same as fatality risk to right front passengers to within 1 per cent; this was so for crashes in all directions and for frontal crashes. Fatality risk in rear seats was (26 ± 2) per cent lower than in front seats, and lower in center compared to outboard seats by (22 ± 4) per cent for front seats and (15 ± 4) per cent for rear seats. The center rear seat was associated with the lowest fatality risk. (*Am J Public Health* 1988; 78: 1456–1458.)

Introduction

A car driver's physical environment differs from that of a front seat passenger because of the location of vehicle controls, especially the steering wheel. However, it is not known whether the driver or right front seat passenger is at greater injury risk. One of two analyses[1] found drivers less likely to be injured, while the other analysis found no difference. There are even larger differences in physical environment between the front and rear seats, and lower fatality risk has been associated clearly with the rear seat positions.[2]

A major problem in comparing safety of the driver and right front seating positions is that they tend to be used by occupants with systematically different probabilities of fatality in crashes. For example, right front passengers are disproportionately female, and females are more likely than males to be killed in crashes of the same severity.[3] Similarly, fatality risk depends strongly on occupant age.[3] The potentially confounding effects of sex and age are in fact larger than any anticipated fatality risk differences attributable solely to seating position. Attempts to remove such large confounding effects through multivariate analyses are unlikely to be satisfactory when the effect of interest is small compared to the confounding effects.

We therefore used a matched-pair analysis[4] to estimate the risk in various seating positions relative to that in the driver seat (hereafter referred to as the relative risk), using data from the Fatal Accident Reporting System [FARS][5] for 1975 through 1985.

Method

Data were selected from the Fatal Accident Reporting System (FARS), a computerized data file maintained by the National Highway Traffic Safety Administration containing detailed information on all traffic crashes occurring in the United States since January 1, 1975 in which anyone was killed.[5] Adult (age ≥ 16 years) occupants of passenger cars were selected from the FARS data for 1975 through 1985; occupants coded as using any type of restraint were excluded. Those with restraint use coded as unknown (about 30 per cent of all the data) were retained, even though about 4 per cent of these probably used belts.

The relative risk of fatality, R, to a passenger in some specific seat to that of the driver in the same car was estimated by confining comparisons to those cases in which driver and passenger were the same sex, the same age (within three years), and neither was coded as restrained.[4] That is

$$R = \frac{N_P}{N_D}, \qquad (1)$$

where N_P and N_D are the numbers of passengers and drivers killed, respectively, in cars containing drivers and passengers (in the specific seat being investigated).

By assuming that the observed fatality frequencies arise from a Poisson process, we estimated the standard error, ΔR, in R as

$$\Delta R = R\sqrt{1/N_D + 1/N_P}. \qquad (2)$$

Because the interpretation of R as relative fatality risk is essentially assumption free, there was no need to include an extra term reflecting possible systematic, or instrument-like, errors, as was done in earlier work (see, for example, equation 14 of Ref. 6 or equation 5 of Ref. 4).

Results

Figure 1 shows that for cars containing a driver and a

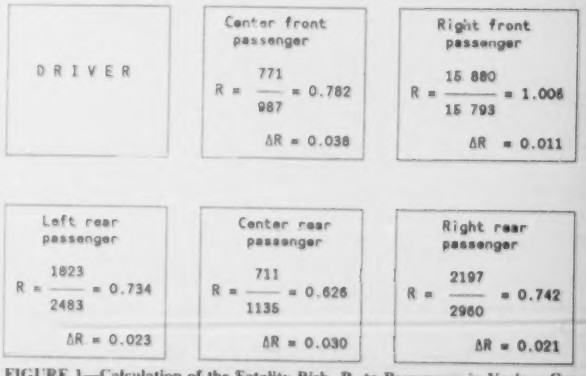

FIGURE 1—Calculation of the Fatality Risk, R, to Passengers in Various Car Seat Positions Relative to That for the Driver in the Same Car
NOTE: Numerators and denominators are, respectively, the numbers of passengers and drivers of the same sex with ages the same to within three years, and not coded as using any restraint system

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TABLE 1—Numbers of Passenger and Driver Fatalities by Principal Impact Point in Clock Points from 1 O'clock through 12 O'clock (12 o'clock may be interpreted approximately to indicate head-on impact; the five car passenger seating positions are as in a bird's eye view of the car, shown in Figure 1).

Principal impact pt

DRIVER

	N_P	N_D	R	δR	N_P	N_D	R	δR
1	30	33	0.91	0.23	1198	550	2.18	0.11
2	32	21	1.52	0.43	638	385	1.66	0.11
3	116	61	1.90	0.30	3095	1126	2.74	0.10
4	4	7	0.57	0.36	150	89	1.69	0.23
5	4	5	0.80	0.54	108	82	1.32	0.19
6	7	6	1.17	0.65	205	205	1.00	0.10
7	5	4	1.25	0.84	68	90	0.76	0.12
8	6	8	0.75	0.41	84	117	0.72	0.10
9	51	148	0.34	0.06	862	2244	0.38	0.02
10	11	25	0.44	0.16	215	442	0.49	0.04
11	34	61	0.56	0.12	546	1145	0.48	0.02
12	244	295	0.83	0.07	5334	5401	0.99	0.02

1	73	89	0.82	0.13	30	31	0.97	0.25	114	115	0.99	0.13
2	42	51	0.82	0.17	24	26	0.92	0.26	61	58	1.05	0.19
3	130	125	1.04	0.13	79	69	1.14	0.19	446	167	2.67	0.24
4	27	11	2.45	0.88	13	5	2.60	1.37	83	12	6.92	2.14
5	9	11	0.82	0.37	12	4	3.00	1.73	52	12	4.33	1.39
6	55	25	2.20	0.53	24	8	3.00	1.22	63	35	1.80	0.38
7	29	15	1.93	0.61	7	7	1.00	0.53	18	16	1.13	0.39
8	76	19	4.00	1.03	11	8	1.38	0.64	27	19	1.42	0.43
9	287	350	0.82	0.07	57	140	0.41	0.06	164	404	0.41	0.04
10	41	62	0.66	0.13	7	18	0.39	0.17	30	87	0.34	0.07
11	131	208	0.63	0.07	32	92	0.35	0.07	77	225	0.34	0.05
12	581	920	0.63	0.03	245	421	0.58	0.05	661	1084	0.61	0.03

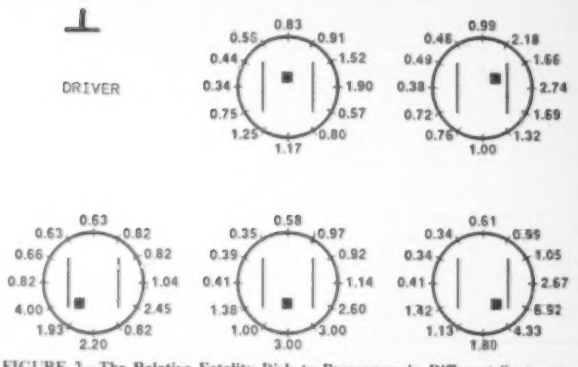

FIGURE 2—The Relative Fatality Risk to Passengers in Different Seats as a Function of Principal Impact Point
NOTE: The values shown are calculated in Table 1, which also provides the raw data

right front passenger, there were 15,880 right front passenger fatalities compared to 15,793 driver fatalities, for a right front passenger relative fatality risk of 1.006 ± 0.011. The center front seat R = 0.78 ± 0.04 indicates that this position was associated with a (22 ± 4) per cent lower fatality risk than the front outboard positions. This substantial difference probably reflects protective effects from the outboard occupants in the form of impact cushioning and ejection prevention.

The outboard rear seats were associated with lower fatality risk than the corresponding front seats. The relative risks in the two outboard positions did not differ, and may be combined to give a composite value for rear outboard seats of R = 0.739 ± 0.015. That is, for unrestrained occupants in outboard seating positions, rear seats were associated with a fatality risk (26.1 ± 1.5) per cent lower than for front seats. Comparing the center rear seat relative fatality risk of R = 0.626 ± 0.030 to R = 0.739 ± 0.015 for the rear outboard seats reveals that the fatality risk in the center rear seat was (15 ± 4) per cent less than in the rear outboard seats.

Table 1 shows data in the five passenger seating positions versus principal impact point, as coded in FARS in terms of clock points from 1 to 12. Although damage at, say, clock point 12 o'clock, may arise from the front of the vehicle being struck at an oblique angle, a principal impact point at 12 o'clock typically indicates frontal crashes. Table 1 shows the number of passengers killed for crashes with the indicated principal impact point divided by the number of driver fatalities for the same principal impact point. A high value of R may reflect an increased passenger fatality risk or a decreased driver fatality risk, or some combination of these. The seating locations follow the same representations as in Figure 1—a bird's eye view of the vehicle from above. The data were more concentrated at 3, 6, 9 and 12 o'clock than at adjacent clock points; these specific values will be discussed below. The values of R in Table 1 are shown schematically in Figure 2.

For principal impact point 12 o'clock, which may be interpreted approximately as indicating head-on impact, the value of R for right front passengers was R = 0.988 ± 0.019. Thus the similarity of fatality risk to drivers and right front passengers applies also to the frontal case.

In frontal crashes the advantage of rear compared to corresponding front seating positions was larger than for all crash directions combined. By comparing rear to front fatality risk in the left, center, and right seating positions, we found reduced risks in the rear of 37, 30, and 38 per cent, respectively. In comparing the results for frontal crashes only (12 o'clock in Table 1 and Figure 2) to those for crashes in all directions (Figure 1), we found that the frontal crashes were associated with a larger fatality reduction for rear seat compared to front seat occupancy, but a diminished reduction was associated with center compared to outboard occupancy. Both these findings are consistent with intuitive understanding of crash processes. The finding that there was still a reduced fatality risk in the center front (R = 0.83 compared to R = 1 and 0.99 in the outboard positions) and center rear (R = 0.58 compared to 0.63 and 0.61 in the outboard positions) may reflect some sideways occupant motion even for principal impact point at 12 o'clock, with the center occupant receiving cushioning and ejection protection from the outboard occupants.

For principal impact point at 6 o'clock, generally a rear impact, occupants in rear seats faced considerably higher fatality risk than those in front seats. Combining data for all three rear seats (Table 1) gives R = 142/68 = 2.09 ± 0.31; so occupants in rear seats were (109 ± 31) per cent more likely to be killed than were drivers. Despite this large difference, Table 1 shows that only a small fraction of rear seat occupant fatalities were due to rear impact. This is because rear impact crashes tend to be of low severity, even though there are many such crashes; typically, rear impact results from "fender-bender" types of two-car crashes in urban traffic, whereas a typical fatal crash is a single car impacting a fixed object, such as a tree, at high speed.

For principal impact point at 3 o'clock, passengers seated on the right side of the vehicle were at much greater risk than the driver. The relative risk was similar for front seat passengers (R = 2.74 ± 0.10) and rear seat passengers (R = 2.67 ± 0.24), for a weighted average value of 2.73 ± 0.09. Independent of whether they were seated in front or rear, occupants on the left side of the vehicle were subject to similar fatality risk (R = 1.04 ± 0.13 for the left rear passenger compared to R = 1 for the driver). Center occupants, whether

in front or rear, were more at risk than drivers in right impacts. Being in a right outboard position compared to being in a left outboard position increases fatality risk by (173 ± 9) per cent.

For principal impact point at 9 o'clock, the nominal indication was that the left rear passenger was at lower risk than the driver (R = 0.82 ± 0.7). If clock points 8 (with an anomalously high value of R), 9 and 10 are combined, the result becomes R = 0.94 ± 0.06, which indicates that the fatality risks to driver and left rear passenger do not differ in left impact crashes. The risk in left impact was substantially less for passengers in right seats than in left seats, but as in the corresponding situation for right impact, there was no indication that in a left side impact there was any difference in fatality risk between the right front (R = 0.38 ± 0.02) and right rear (R = 0.41 ± 0.04) seating positions. In right side impacts, 3,095 right front passengers were killed compared to 2,244 drivers killed in left side impacts, suggesting 38 per cent more impacts of high severity from the right than from the left.

For left side impacts, the driver relative to passenger fatality risk (that is, 1/R) was 2.60 ± 0.10 for front seat passengers and 2.46 ± 0.22 for rear seat passengers, for a weighted average of 2.58 ± 0.09; that is, the near side occupant was (158 ± 9) per cent more likely to be killed. This may be compared to the corresponding result of (173 ± 9) per cent for the right impact case. The quantitative agreement for these physically symmetric cases, and the general qualitative agreement found between many of the above results and physical intuition, adds credibility to the method and to confidence that the quantities used were coded with satisfactory reliability in the FARS data.

Discussion

The finding that the fatality risk to right front passengers was the same as that to drivers, to within about 1 per cent, was surprising in view of the substantial differences between the physical environment of drivers and right front passengers.[7,8] The similarity of fatality risk applies also to head-on crashes. Results apply only to fatality risk, and may not apply to other levels of injury. Large numbers of data permitted more precise estimates than have previously been available, and the close agreement of estimated risks for physically symmetric crashes lends credibility to the methods and data used. The estimates reflect the mix of cars and crash types included in the 1975 through 1985 FARS data. Another study[9] using the same method found that, for cars of model year 1966 and earlier, the right front passenger fatality risk exceeded that of the driver by (9.6 ± 6.5) per cent, whereas for cars of model year 1968 and later, the right front passenger fatality risk was (2.8 ± 2.0) per cent less than that for the driver.

Although all the results presented were for unrestrained occupants, it was possible to estimate relative fatality risk to restrained occupants in outboard seats using published estimates of restraint system effectiveness in preventing fatalities in outboard seats.[10,11] Taking lap/shoulder belt effectiveness[10] as (43 ± 3) per cent gives the relative fatality risk for occupants (necessarily outboard) using such systems as 0.57 ± 0.03. Taking lap-only belt effectiveness[11] for outboard rear seats as (18 ± 9) per cent gives a relative fatality risk to lap-belted outboard rear occupants of 0.739 × 0.82 = 0.61 ± 0.07, which, when the error limits are taken into account, is not different from the value for outboard front occupants. Hence, when all occupants use the most commonly provided restraint systems, no difference is indicated in the fatality risk to front or rear outboard occupants.

Examining relative fatality risk according to direction of impact showed that occupants closer to the point of impact were at much greater risk of fatality than those further from the point of impact. It must be remembered that these are relative risks and not absolute risks. For example, rear seats have lower overall fatality risk notwithstanding their high relative fatality risk in rear crashes because such crashes are so small a fraction of all fatal crashes (see Table 1).

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Sunlight and Other Risk Factors for Cataracts: An Epidemiologic Study

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Abstract: A case control study was conducted in North Carolina to explore the relation between individual exposure to sunlight and the risk of cataracts. One hundred thirteen cases and 161 controls aged 40–69 at diagnosis were studied. Sunlight exposure was inferred from interview data on residency and time spent in the sun, combined with solar radiation data from the National Climatic Data Center. Sunlight exposure was very slightly related to all types of opacities combined. Although the numbers of cases with each type of opacity was small, the risk of cataracts was slightly increased in medium and high exposure categories for persons having cortical or posterior subcapsular opacities only, but not nuclear sclerotic changes. Persons with dark brown or hazel eyes are at increased risk. An unexpected finding was that persons who reported using tranquilizers for six months were at increased risk. (*Am J Public Health* 1988; 78: 1459–1462).

Introduction

Throughout the world, cataracts are most prevalent in areas where the amount of annual and daily sunlight is high[1–3] ambient temperatures are warm,[4] and the latitude is low.[3] These areas receive more solar radiation, which includes ultraviolet and infrared radiation and visible light.

Environmental measures of solar radiation exposure have been used in previous studies as surrogates for individual levels of sunlight exposure. In addition, these studies have relied on a single estimate of exposure, based either on current[2,4] or usual residence.[5] Such measures cannot correctly estimate the amount of sun exposure an individual receives, since lifestyle factors influence the amount of time spent in the sun within a geographic area, and since sun exposures change over the subject's lifetime.

To investigate this question further, we conducted a case-control study in North Carolina which utilized a detailed estimate of sunlight exposure for each study subject.

Methods

Study Population

All study subjects were selected from the patients of a private ophthalmology practice in Asheboro, North Carolina. During their first visit to the practice, each patient received a full ocular examination including both external and internal examination of the eye. Direct visualization of the lens using a Zeiss slit lamp was performed after pupil dilation (using 2 1/2 per cent Neosynephrine and 1 per cent Mydracil). All were examined by a single ophthalmologist (ASL).

Cases and controls were chosen by reviewing medical records of all patients aged 40–69 years old at their first examination (1980–83) (n = 864). Ineligible patients included those who were blind (before detection of an opacity), deceased, or did not have a working telephone number listed in their medical record. Cases who had cataracts removed by another physician were excluded because characteristics of the lens were not available. Controls were excluded if they had photokeratitis (a corneal burn due to high solar exposure).

Patients were considered to be cases if their chart indicated the presence of cortical, posterior subcapsular or nuclear sclerotic changes of the lens in either one or both eyes. Ophthalmic information abstracted from the cases' records included: type of cataract, severity of the opacity, number of opacities, eyes involved, surgery, visual acuity and comorbid ocular conditions. Cases were included regardless of visual acuity; and diagnoses were not validated in any other way. There were 133 cases identified in the population, 15.4 per cent of persons aged 40–69 years. Cataract type was classified separately for subjects with one type of opacity and multiple types of opacities. Patients with either one opacity in one eye or the same type of opacity in both eyes were classified in the single opacity type group. Multiple opacities either in one or both eyes were classified according to the combination.

Controls were selected from the remaining patients in the targeted age group. Their records showed no indication of opacification of the lens of either eye during the time of the study. Two controls similar to the cases with respect to current age (± two years) and sex were chosen, when available, otherwise one control was selected. One hundred and ninety-three controls were enrolled into the study.

Eighty-seven per cent (n = 283) of eligible subjects responded to the telephone interview, which gathered historical information on exposure to natural sunlight, demographic characteristics, ophthalmic factors, use of selected medications with photosensitizing properties, medical conditions, selected occupations, smoking, alcohol use, and caffeine intake. Few study subjects were Black (n = 9), so the analysis was restricted to the 113 White cases and 161 White controls.

Sun Exposure Measures

The variable for sun exposure was derived from various sources of questionnaire information. A complete residential history, from place of birth to current residence, was obtained for each study subject. For all residences of one year or more, the subject was asked the state and number of years of residency. The respondent was also asked to estimate the amount of time spent in the sun during those years. Subjects were given a choice of three responses (much time in the sun, a moderate amount, or a little time in the sun). Up to eight residences were recorded for each participant.

Mean total solar radiation values for each state were obtained from the National Climatic Data Center[6] which has

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TABLE 1—Characteristics of the Population

Characteristics	Cases Number (%)	Controls Number (%)
Sex		
Male	45 (40)	69 (43)
Female	68 (60)	92 (57)
Age*		
40–49	10 (9)	18 (11)
50–59	26 (23)	48 (30)
60 +	77 (68)	95 (59)
Education		
Less than High School	50 (44)	65 (40)
High School Graduate	46 (41)	57 (35)
More than High School	17 (15)	39 (24)
Diabetes		
Yes	18 (16)	21 (13)
No	95 (84)	140 (87)
Lifelong Residents of NC		
Yes	74 (65)	107 (66)
No	39 (35)	54 (34)

*Age (years) in 1983.

routinely measured daily direct and diffuse solar radiation on a horizontal surface at weather reporting stations in each state over a 47-year period (1915–62), although not all weather stations had been in operation for the full 47 years. An average solar value was calculated for states with more than one reporting station. Intensity of the solar radiation is measured in Langley units which equal one gram calorie per square centimeter.

An individual's average annual sun exposure was a weighted average of lifetime sun exposure. The radiation intensity of each state was multiplied by the duration of time spent in that state times the reported amount of time spent in the sun (assigned a proportionate value of 80 per cent for much time in the sun, 50 per cent for a moderate amount of time, 20 per cent for a little time in the sun). This product was summed over the various states of residence and divided by the respondent's age.

Statistical Methods

Unconditional logistic regression models were fitted with lens disease as the dependent variable and with the matching factors (age and sex) included in all models. The exposure variable, annual sunlight exposure, was included in various models as a continuous variable, dichotomized at the median value of the distribution or grouped into three categories (low, medium, and high). Screening procedures were used to identify potential confounders. Odds ratios and 95% confidence intervals were estimated from the logistic models. Separate models were fitted for each type of lens opacity with enough cases.

Results

As shown in Table 1, approximately 60 per cent of the study population was female. The majority of the cases and controls were elderly, non-diabetic lifelong residents of North Carolina.

Seventy-four per cent of all cataract cases had a single opacity type, as seen in Table 2. Nuclear sclerotic changes and posterior subcapsular changes predominated. Most cataract cases had opacities in both eyes, and few had greatly impaired visual acuity.

The mean annual sun exposure score was 212.40 Langley/year (SE=7.63) for cases and 207.62 Langley/year (SE=5.82) for controls. The average number of hours spent in the sun per day as an adult was 5.0 hours/day for cases and 4.9 hours/day for controls.

TABLE 2—Ophthalmic Characteristics of the Cataract Cases

Characteristics	Number (%)
Cataract type	
Single opacity	
Cortical	9 (8)
Posterior Subcapsular	13 (12)
Nuclear Sclerotic	56 (50)
Other	5 (4)
Multiple opacities	
Cortical & Posterior subcapsular	4 (3)
Cortical & Nuclear Sclerotic	3 (3)
Posterior Subcapsular & Nuclear Sclerotic	22 (20)
Number of Eyes Involved	
Unilateral	29 (26)
Bilateral	84 (74)
Number of Opacities—Right Eye	
0	14 (12)
1	83 (74)
2	16 (14)
Visual Acuity—Worst Eye	
20/25 or better	43 (39)
20/30—20/40	21 (19)
20/50—20/90	23 (21)
20/100 or worse	24 (21)

As presented in Table 3, for all types of opacities combined, the odds ratios were 1.06 for moderate exposure and 1.12 for high exposure compared with low exposure to sunlight. The number of cases with each type of opacity was small and the odds ratios correspondingly unstable. Sunlight exposure appeared to relate to cortical and posterior subcapsular opacity but not to nuclear sclerotic lens changes. These small changes in risk were also apparent when the sun exposure data were treated as continuous or dichotomized at the median.

Additional variables were examined as potential risk factors for cataract, as seen as Table 4. The use of tranquilizers was positively associated with the risk of lens opacities (OR=2.20, 95% CI=1.10, 4.39), with 15 per cent of the total study population reporting having used tranquilizers for at least six months. The type of tranquilizers used and frequency and duration of their use was not ascertained during the interview. Persons with hazel and brown eyes had higher risk than persons with green, blue and grey eyes. Persons who reported a family history of cataracts had a slightly increased

TABLE 3—Odds Ratios and 95% Confidence Intervals for the Relationship between Sunlight Exposure and the Risk of Cataracts

Cataract Type	Low Exposure	Medium Exposure	High Exposure
All Cataract Cases	1.00	1.06 (0.77, 1.45)	1.12 (0.56, 1.98)
Cortical Only	1.00	1.23 (0.51, 2.98)	1.53 (0.21, 7.19)
Posterior Subcapsular Only	1.00	1.23 (0.59, 2.59)	1.52 (0.28, 5.44)
Nuclear Sclerotic Only	1.00	0.87 (0.77, 1.31)	0.79 (0.39, 1.96)
PSC & NS*	1.00	1.16 (0.65, 2.08)	1.36 (0.36, 3.72)

*Posterior subcapsular and nuclear sclerotic opacities
95% CI in parentheses.

TABLE 4—The Relationship between Cataracts and Other Risk Factors

Risk Factors	Cases		Controls			
	Exposed	Unexposed	Exposed	Unexposed	Odds Ratio*	95% CI
Tranquilizers	24	87	17	140	2.20	(1.10, 4.39)
Dark Eyes	51	60	58	102	1.60	(0.96, 2.65)
Diabetes	18	95	21	140	1.20	(0.60, 2.39)
Physical Injury to Eye	14	98	21	140	0.99	(0.47, 2.14)
Family History of Cataracts	35	77	44	117	1.39	(0.80, 2.41)
Sunglasses	48	63	69	92	1.04	(0.65, 1.77)
Smoking	60	52	83	78	1.28	(0.74, 2.22)
Outdoor Work	19	91	33	128	0.86	(0.57, 1.75)

*Odds ratio adjusted for age and sex.

risk of lens disease. Diabetes, sunglass use, eye injury, and current smoking habit had odds ratios close to the null. A history of outdoor work which may involve higher sunlight exposure was not related to an increased risk of lens disease in either men or women.

Discussion

To our knowledge, this is the first case-control study conducted to estimate the risk of lens disease according to exposure to natural sunlight. Cases had a slightly higher mean annual sun exposure score than controls but reported an equal number of hours spent in the sun as adults. Our study showed a very weak increase in the overall risk of opacities as lifetime exposure to the sun increased. An increasing trend was seen for risk of cortical or posterior subcapsular cataracts but not nuclear sclerotic cataracts, as well as cases having a combination of opacity types, but these estimates are imprecise because of the small number of cases of each type. The very modest increase in risk we observed is smaller than previously reported from ecologic studies. Hiller, *et al.* [5] using HANES data, reported a risk ratio of 1.58 for Tucson, Arizona compared with Albany, New York. Studies done in Nepal[4] found a prevalence ratio of 2.6 for villages with 12 hours of sunlight compared with villages with 7–9 hours of sunlight. Similar patterns have been reported within Australia[2,7] and the United States.[3] Another study found more cataracts in Manila than in Tampa, Florida and Rochester, New York.[8] Additional analysis of the HANES data found exposure to ultraviolet-B radiation to be significantly associated with cortical changes (RR=3.6, n=55).[11] A positive association was found for nuclear sclerotic cataracts (RR=1.4, n=104) and a negative one for posterior subcapsular changes (RR=0.3, n=18). One likely reason for lower risk levels in these data is that 65 per cent of the study population were lifelong residents of North Carolina and therefore the range of exposure scores was narrower than in the nationwide sample.

Our risk estimates may also be biased toward the null because of limitations in the characterization of lifetime exposure. Solar radiation measurements for the state of residence were quite accurate, but the duration of residence and amount of time spent in the sun could have been incorrectly remembered or interpreted differently among study participants. No direct validation of this exposure classification scheme was possible, but we did note that the average annual sun exposure scores were higher for persons working outdoors than indoors (270.13 Langley/yr, SE=7.1 versus 194.37 Langley/yr, SE=5.0), and persons reporting greater than the median number of hours in the sun had higher scores than persons spending less time outdoors (245.35 Langley/yr, SE=5.6 versus 168.40 Langley/yr SE=5.6).

Our estimates of the relative risk of each of the individual cataract types require cautious interpretation because of the study size. We restricted our study population to ages 40–69 at age of diagnosis to separate the etiologic effects of solar radiation and aging and to improve recall, but there are fewer lens changes occurring in this younger group. A larger study would lend more confidence to our findings.

If the modest gradient in risk with solar radiation in this study and the marked gradient seen in population studies is real, ultraviolet radiation may be the relevant part of the solar spectrum in the etiology of cataracts. This may be of particular importance for ocular health in light of recent reports of future amounts of ultraviolet radiation in our atmosphere from depletion of the ozone layer. Ultraviolet radiation has been used to create cataractous changes in animal lenses.[16] Both UV-A and UV-B are capable of producing lenticular damage.[17] Cataracts are produced experimentally using either a single high dose of ultraviolet radiation or multiple exposures to lower intensity. One may hypothesize that repeated insults of low intensity radiation over the course of a persons lifetime leads to cataracts. Damage may be done episodically, when sun exposure is of high intensity or long duration, with accumulations of aberrations ultimately leading to opacification. The slow progression of lens disease is consistent with this notion of accumulated damage in the lens.

Other risk factors for cataracts were found in our study. The apparent effect of tranquilizers was an unanticipated finding. Bartholomew and others[8] reported no association between the use of major tranquilizers and cataract. In our data, the tranquilizers used were probably sleeping pills or other barbiturates and minor tranquilizers such as diazepam, rather than "major tranquilizers". More detailed data on tranquilizers used should be included in future studies.

Our finding of increased risk among dark-eyed subjects is consistent with reports of high prevalence rates of cataracts in areas populated with dark-eye persons, e.g., India,[10] Nepal,[4] and Aborigines in Australia.[2] Black race was also reported as a significant risk factor for cataracts in the HANES data, but our study population had too few Blacks to study this factor.[5] While melanin protects against the carcinogenic effects of ultraviolet radiation to the skin, melanin in the iris may absorb solar radiation and provide higher amounts of exposure to the lens of the eye.[12] This finding needs replication in other population studies. Our study did not find factors such as eye injury[13] or diabetes[14,15] to be related to lens changes, although they have been

reported as risk factors in other studies. Smoking was evaluated for the first time and found not to be related to cataract risk. The positive associations between lens changes and eye color and tranquilizer use may stand up since a large number of comparisons were made and our numbers are small.

Other studies of lens disease are needed to confirm our findings and those of previously reported geographic studies. Residential history coupled with objective measurements of sun exposure, as attempted in this study, is necessary. Larger populations of cases and controls with substantial variation in sun exposure are also needed.

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NCHS Report Analyzes Prevalence of Chronic Conditions in US

An analysis of the prevalence of selected chronic conditions in US citizens was recently published by the National Center for Health Statistics, using data from its annual National Health Interview Survey 1983-85.

The report ranks the top 10 chronic conditions for the total population, including breakdowns by age, sex, and race. It also shows the 15 most prevalent chronic conditions and the percentage of persons ever hospitalized due to them; chronic conditions were also ranked according to limitation of activity.

There was little difference between sexes in ranking of the top 10 chronic diseases; nine out of 10 chronic conditions ranked equally high.

- Sinusitis was the most prevalent chronic condition (31.2 million reported conditions annually); arthritis ranked second (30.3 million conditions); high blood pressure was third (28.6 million).
- Heart disease and diabetes were the chronic diseases most likely to require hospitalization (42.6 and 32.7 per cent, respectively).
- Chronic conditions causing highest limitation of activity were mental retardation (85.6 percent), followed by multiple sclerosis (76.8percent) and cancer of the lung and bronchus (68.2 per cent).
- Those under age 18 reported more respiratory conditions (hay fever, bronchitis, sinusitis and asthma) while the 65+ age group reported more cataracts, tinnitus, diabetes and hardening of the arteries.
- Among Blacks, high blood pressure was the most prevalent chronic condition reported (146.3 per 1000).

The top 10 chronic conditions in prevalence were very different from those with high levels of hospitalization and activity limitation. Less than 10 per cent of people with the three most prevalent conditions were hospitalized; sinusitis, the most prevalent chronic condition, caused limitations in less than .005 per cent of those reporting the condition.

Copies of the 14-page report, *Prevalence of Selected Chronic Conditions, United States 1983-85*, Advance Data Report No. 155, are available free from: National Center for Health Statistics, USPHS, DHHS, 3700 East West Highway, Room 1-57, Hyattsville, MD 20782.

Public Health Briefs

Screening Prospective Blood Donors for AIDS Risk Factors: Will Sufficient Donors Be Found?

DAVID I. GREGORIO, PHD, MS, AND JEANNE V. LINDEN, MD, MPH

Abstract: Using data from various sources–we estimate that 14 to 19 per cent of American males 17–75 have personal histories that place them at high risk of transmitting the HIV infection while an additional 2 per cent of adult females may be similarly affected. Because roughly one fourth of either group may already be unsuited to give blood, we estimate that 10–14 per cent of adult males, and 1 per cent of females would be specifically deferred from giving blood because of personal histories of high-risk behavior. Local adjustments in the assumptions underlying these estimates are needed to apply them to given communities. (*Am J Public Health* 1988; 78:1468–1471.)

Introduction

Since 1983, blood collection centers in the United States have advised persons with AIDS (acquired immunodeficiency syndrome) or a history of behavior placing them at high risk of HIV (human immunodeficiency virus) infection to avoid donating blood.[1] Limited evidence suggests that voluntary self-deferral guidelines have produced the desired effect.[2–6]

Wykoff and Halsey[4] found that the blood donation rate by persons who were subsequently diagnosed with AIDS or Aids-Related Complex was reduced about 50 per cent following the implementation of self-deferral guidelines by blood collection centers in the New Orleans area. Kalish, *et al*,[5] similarly observed that the frequency of seropositivity detected among Connecticut blood donors fell approximately 75 per cent. Pindyck, *et al*,[6] reported that implementation of self-deferral guidelines by the Greater New York Blood Program (GNYBP) resulted in significantly fewer male donors appearing at blood collection sites, along with more potential donors being rejected because of personal histories of high-risk behavior uncovered through pre-donation health history screening, and a greater number of confidential requests from accepted donors that their blood not be used for transfusions.

Those results, along with our previous finding that a sizable number of adult Americans already may be unsuited to give blood,* prompted interest in estimating what proportion of our population that is otherwise eligible would be deferred as blood donors because of personal histories of high-risk behavior.

Methods

The overall proportion of eligible blood donors in the population is defined as adults, 17–75 years, who are not included within one or more of the American Red Cross/American Association of Blood Banks (ARC/AABB) donor deferral categories.[7] Of the many criteria currently used to screen prospective blood donors, the 16 most common reasons for deferring donations are summarized in the Appendix. As used in this exercise, they are specific to the American Red Cross Blood Services, Connecticut Region, although similar standards are used elsewhere.

Among the donor deferral guidelines cited, the first eight address matters of donor safety; the remainder pertain to the safety of blood products for use in transfusions. Of the latter eight criteria, three specifically address behaviors that increase a person's risk of HIV infection: histories of homosexual activity or prostitution as well as sexual encounters with prostitutes or other high-risk individuals.

For each of the 16 donor deferral categories summarized in the Appendix, established national data sources were used to estimate the proportions of our adult male and female populations possessing traits that would render them unsuited to give blood.[8–15] Whenever nationwide prevalence data were scarce, estimates were based on research findings by recognized experts.[16–28] As necessary, a range of prevalence estimates was considered to reflect "best" and "worst" case scenarios for estimating overall donor eligibility. All age- and sex-specific calculations were derived using the 1980 US population census.[29]

Since prospective donors could be unsuited for more than a single reason, the overlap among various traits was corrected according to the following rules: Whenever *a priori* associations among criteria could be assumed (e.g., intravenous (IV) drug use among female prostitutes), the unduplicated prevalence of either factor was estimated to equal the reported prevalence of the specific traits, minus the estimated overlap with other categories. When *a priori* associations could not be put forth (e.g., a history of heart disease among female prostitutes), we conservatively reasoned that since the probability of independent events occurring jointly is the product of their independent probabilities, such overlap would be proportional to the respective size of the specific deferral categories. Accordingly, the proportion of the population exhibiting one or more behaviors placing them at risk of transmitting HIV infections (and thus deferrable as blood donors) can be estimated as one minus the product of the

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TABLE 1—Estimated Proportion and Number of the US Adult Population Unsuited to be Blood Donors by Sex and Criteria for Exclusion

Criteria for Donor Deferral†	Estimated Proportion of US Population	
	% Males	% Females
Donor safety (criteria 1–8)	18–22	39–41
Blood product safety (criteria 9–16)	20–25	8
Miscellaneous risk factors (criteria 9–13)	7	8
AIDS risk factors (criteria 14–16)	14–19	2
Homosexual activity	10–15	0
Sex partners at high risk	6	3
Prostitution/sex with prostitutes	<1	<1
Total Proportion of Population Deferred†††	34–41	42–45
Total Population, 17–75 yrs old††	80,331,000	84,404,000

†Definitions of donor deferral criteria and groupings of deferral categories from Table 1.
††Source: US Bureau of the Census.[29]
†††Total of deferred population may exceed category entries because of the overlap among traits.

probabilities that persons are not included within one of the three AIDS risk factor deferral categories. Similarly, the unique proportion of the population deferred as blood donors because of AIDS risk factors, but otherwise suited to give blood (i.e., not already deferrable by criteria 1–13 of the Appendix table), is estimated to equal the proportion of AIDS risk factors among the population, multiplied by the proportion of other conditions that warrant deferral of blood donors.

Results

The estimated prevalence of traits that would require the deferral of blood donors are presented for both males and females in Table 1. Regarding criteria that pertain to blood donor safety (i.e., categories 1–8 of the Appendix), we estimate that 18 to 22 per cent of adult males and 39–41 per cent of women may be unsuited for giving blood. Among males, the greatest proportion of deferrals would be for reasons of a history of hypertension, having low hematocrit, or being over age. Women, on the other hand, would most likely be deferred for reasons of age, inadequate weight, or low hematocrit.

Between 20–25 per cent of adult men may be affected by conditions which threaten the safety of donated blood products. The prevalence of such traits among adult women, however, is not thought to exceed 8 per cent of that group's total population. The disparity stems from a greater prevalence of traits related to AIDS risk factors among males (i.e., categories 14–16 of the Appendix).

We conservatively estimate that 10–15 per cent of adult males would be unsuited as blood donors for reason of a homosexual experience since 1977. Sex partners of individuals who are at high risk of AIDS infections was estimated to affect 6 per cent of the adult male population and perhaps one-half of that amount among adult females. We further estimate that a history since 1977 of prostitution and/or a sexual encounter with prostitutes would affect considerably smaller proportions of the population.

Considerable overlap of AIDS risk factors among the population is known to exist. Therefore, the total portion of the population to be deferred as blood donors because of AIDS risk factors would be somewhat smaller than the simple sum of the estimated prevalences of particular AIDS risk factors among the population. For example, it is thought that one-half of all prostitutes, and those persons having had a sexual encounter with prostitutes, use IV drugs.[25] As many as one-third of male IV drug users are believed to have had histories of homosexual activity,[26] while a similar proportion of female drug users are believed to have engaged in prostitution.[27] Perhaps as many as one-half of women whose sex partners are from high-risk groups have themselves used IV drugs or engaged in prostitution, while one-fourth of males with sex partners from other high-risk groups have a history of IV drug use.[28] Taking these factors together, we estimate that 14–19 per cent of American males between the ages of 17 and 75, and roughly 2 per cent of adult females have personal histories that increase their risk of transmitting the HIV infection.

We calculated that donor deferral criteria unrelated to AIDS risk factors (i.e., criteria 1–13 of Table 1) could affect roughly one-fourth of adult males, and nearly one-half of females. Reducing estimates of the portions of the male and female populations with AIDS risk factors by such amounts to account for overlap among deferral categories suggests that 10–14 per cent of adult males, but perhaps as few as 1 per cent of females would be specifically deferred from giving blood because of a personal history of high-risk behavior. These amounts represent the unique reduction of our blood donor pool that might be attributed to ongoing efforts to screen individuals for AIDS risk factors.

Discussion

Of the roughly 165 million adults in this country 17–75 years of age, some 64 million persons (37 per cent) may not be suited to give blood for one of the various reasons cited in the above analysis. Of that number, we estimate that approximately 90 per cent would remain ineligible in the absence of AIDS risk factor screening. Individuals unsuited for giving blood specifically for reasons of being at greatest risk of transmitting HIV infections are thought to number between 8 and 11 million persons.

We caution, however, that data used here for national estimates may be unreliable and that the prevalence of specific traits will vary markedly from place to place.[30] Local adjustments in the assumptions underlying these estimates are needed to make them applicable to given communities.

Nevertheless, we are confident that the pool of suitable blood donors in this country may equal 100 million persons. Presently, only about 8 million adults (roughly 8 per cent of the eligible donor pool) donate blood during any year,[16] and surveys suggest that only one-half of Americans give blood during their lifetime.[18] Hence, there are many medically suitable persons who could be recruited as donors to supply blood products for the health care needs of this nation. Until the course of the HIV infection is better understood, persons at risk of transmitting HIV infections should not be permitted to donate blood. The adequacy of our nation's blood supply should not be adversely affected by such a practice.

While autologous transfusion is the simplest strategy to protect individuals from transfusion-related infections, a nationwide blood collection policy based solely on this approach is unwise. Most individuals in need of transfusion either suffer from unanticipated traumatic injuries or they do not qualify as blood donors because of ill health.

Expanded educational efforts to encourage self-deferral of high-risk individuals, in conjunction with donor recruitment efforts to retain former donors and effective laboratory testing of all blood products should be sufficient in assuring the safety of our nation's blood supply.

APPENDIX

Principal Criteria Used in Deferring Prospective Blood Donors

Deferral Category	Criteria for Deferral	Estimated Portion of US Population	
		% Males	% Females
Criteria Pertaining to Donor Safety (1–8)			
1. Age	Less than 17 or over 75 years of age unless permitted by physician to age 79)	5	9
2. Hematocrit	Less than 38% for women; 41% for men	5	23
3. Weight	Less than 110 pounds	<1	<1
4. Pregnancy	Less than 2 weeks post-abortion; 6 weeks post normal delivery; and 6 months post-cesarean section	0	3
5. Major Surgery	Less than 6 months post-operative	<1	<1
6. Heart Disease	Evidence of significant coronary artery disease (Coronary artery bypass graft, chest pain, arrhythmias, myocardial infarctions)	4	3
7. Diabetes	Present use of insulin	1	1
8. Hypertension	Systolic level >180 torr; Diastolic >100 torr	5–10	4–8
Criteria Pertaining to Safety of Blood Products (9–13)			
9. Cancer	History of cancer other than local skin cancer or carcinoma-in-situ	1	1
10. Transfusion	Less than 6 months since receipt of blood products	<1	<1
11. Non-A Non-B Hepatitis Testing	Elevated ALT/positive for anti-HBc antibody	4	4
12. Miscellaneous	Medications, travel to malarial/AIDS risk areas, exposure to hepatitis, etc.	<1	<1
13. IV Drug Use	Single episode	2	1
Criteria Pertaining to AIDS Risk Factors (14–16)			
14. Homosexual Activity	Single homosexual act since 1977	10–15	0
15. Sex with a Prostitute/Prostitution	Single episode since 1977	<1	<1
16. Sex Partner at High Risk	Single sexual contact with person in AIDS high-risk groups since 1977	6	3

SOURCES:

1. 1980 US Census[29]
2. National Center for Health Statistics[5]
3. Unpublished data from NHANES-II.[9] Because some people are underweight due to poor nutrition, and thus may be iron deficient, a 50 per cent overlap with low hematocrit is assumed here.
4. 1980 US birthrate[29] with a 25 per cent upward adjustment for pregnancies resulting in abortion. An average of 1.0 births per pregnancy is assumed. A 15 per cent cesarean section rate nationwide is assumed. Calculations of the donor deferral rate at any one time is based on woman-weeks of deferral. Because low hematocrit occurs during and after pregnancy,[17] a 50 per cent overlap with low hematocrit is assumed here.
5. National Center for Health Statistics.[22,23] A 75 per cent overlap with heart disease and transfusions is assumed here.
6. Vital and Health Statistics.[11] One-half of reported disease prevalence is assumed to be significant enough for deferral. A 25 per cent overlap with age and hypertension is assumed here.
7. National Diabetes Data Group.[12] Estimate includes both Type-I and Type-II diabetics taking insulin. A 20 per cent overlap with age is assumed here.
8. NHANES-II[13] for persons 25–74 years of age.
9. National Cancer Institute,[14] Connecticut Tumor Registry and other sources.[19] A 50 per cent 5-year survival for all cases is assumed.[20]
10. Connecticut Region, American Red Cross, and other sources.[18] A total of 3.5 units per recipient is assumed. We assumed 50 per cent of recipients to be long-term survivors and two-thirds of those individuals to be 17–75 years of age.[21] A 66 per cent overlap with age is assumed here.
11. Categories are assumed to be independent.[22,23]
12. Connecticut Region, American Red Cross.[21]
13. New York State Substance Abuse Services 15 estimate for nationwide prevalence of substance abuse, with an incremental adjustment for drug use prior to 1986. Non-intravenous narcotic use is assumed to equal intravenous use. Overlaps of 33 per cent with homosexual activity (among males) and 33 per cent with prostitution (among females) are assumed.[26,27]
14. Kinsey[24] estimated that 10 per cent of males 16 to 55 yers old are more or less exclusively homosexual, while 18 per cent of all males in that group are estimated to have had equal amounts of homosexual and heterosexual activity.
15. Morbidity and Mortality Weekly Report.[25] A 50 per cent overlap with IV drug use is assumed here.[25]
16. One sex partner per female and heterosexual male IV drug user is assumed here[26] with an adjustment of one female partner per 10 bisexual men. Overlaps of 50 per cent with IV drug use and prostitution (among females) and 25 per cent with IV drug use (among males) is assumed here.

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The San Francisco Men's Health Study: Continued Decline in HIV Seroconversion Rates among Homosexual/Bisexual Men

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Abstract: The incidence of infection by the human immunodeficiency virus (HIV) has been monitored since 1984 in an area probability sample of homosexual/bisexual men drawn from a six-kilometer square area of San Francisco where the epidemic of acquired immunodeficiency syndrome (AIDS) has been most severe. Annualized HIV seroconversion rates in previously uninfected cohort members have declined by 88 per cent from 5.9 per cent during the first six months of 1985 to 0.7 per cent during the last six months of 1987. Concurrent declines of approximately 80 per cent in the prevalence of sexual behaviors associated with HIV transmission were also observed in the sample. (*Am J Public Health* 1988; 78: 1472–1474.)

Introduction

Declining sexual transmission of human immunodeficiency virus (HIV) infection among homosexual/bisexual men was first reported from the San Francisco Men's Health Study (SFMHS) in June 1987.[1] From an estimated annual HIV seroconversion rate of 18.4 per cent among seronegative study subjects between 1982 and 1984, the annualized infection rate dropped to 5.9 per cent (95% confidence intervals=2.8, 9.0) during the first six months of 1985 and to 4.2 per cent (95% CI=1.8, 9.3) during the first six months of 1986. These declines in HIV seroconversion rates were attributed to reductions of 60 per cent or more in the prevalence of HIV seronegative homosexual/bisexual men practicing insertive and receptive anal intercourse with multiple partners. Data are now available for HIV seroconversion rates and the prevalence of these high-risk sexual practices in this study population through 1987.

Methods

Sampling, data collection, laboratory, and statistical procedures for the SFMHS have been reported previously.[1,2] The cohort of 1,034 single men 25–54 years of age was obtained by stratified cluster sampling from a six kilometer square area of San Francisco where the AIDS epidemic has been most intense. Four hundred and eleven study subjects were HIV seronegative and homosexual/bisexual on entry into the study and have been examined at six-month intervals since the last half of 1984. The data for this report consist of HIV serostatus, determined by indirect immunofluorescent antibody assay for the first five examination cycles and by a standard enzyme-linked immunosorbent assay (DuPont) for the sixth and seventh cycles with confirmation of seroconversions by immunoblot, and numbers of sexual partners and frequency of anal intercourse determined by interview. The 95 per cent confidence intervals for annualized HIV seroconversion rates were obtained from published tables.[3]

Results

Time Trend of HIV Seroconversion

The data for HIV seroconversions in seven time periods for the SFMHS are shown in Table 1. Since the last report[1] for the period July–December 1986 when the annualized seroconversion rate was 4.2 per cent (95% CI=1.8, 9.3), the rate declined to 1.7 per cent (95% CI=0.3, 4.9) during the first six months of 1987 and to 0.7 per cent (95% CI=.002, 3.8) during the last six months of 1987. During the time the SFMHS cohort has been under observation, the annualized seroconversion rate has dropped by 88 per cent.

In order to evaluate the possibility that the declining HIV seroconversion rate was not due to selection bias, i.e., decreasing participation over time of men practicing high-risk behavior, participation in the SFMHS was evaluated according to risk status. Subjects who were HIV seronegative on entry were classified according to the number of male sexual partners, a major risk factor,[2] reported for the two years prior to the start of the SFMHS. Their rate of participation in subsequent examination cycles was then determined. The data, as shown in Table 2, indicate no relation between risk status on entry and subsequent participation. Furthermore, participation of HIV seronegative study subjects has remained quite constant since the second examination cycle, fluctuating around 80 per cent.

Changing Prevalence of High-Risk Sexual Practices

As already indicated, the sexual practices by which HIV is transmitted between homosexual/bisexual men are insertive anal intercourse by infected men, which disseminates infection, and receptive anal intercourse among uninfected men, by which they acquire infection. The prevalence of these practices with two or more male sexual partners during

TABLE 1—HIV Seroconverters and Seroconversion Rates per Year by Time Period, San Francisco Men's Health Study, July 1982–December 1987

Dates of Seroconversion	Seronegatives Observed	Mean Number of Days Observed	Seroconversions		
			No.	%	(95% CI)
7/82–12/84	615	Not Applicable	204	18.4*	
1/85–6/85	351	212	12	5.9†	(2.8–9.0)
7/85–12/85	311	188	5	3.1†	(1.0–7.4)
1/86–6/86	316	191	7	4.2†	(1.7–9.0)
7/86–12/86	327	185	7	4.2†	(1.8–9.3)
1/87–6/87	301	218	3	1.7†	(0.3–4.9)
7/87–12/87	323	172	1	0.7†	(.002–3.8)

*Estimated as described in Reference 1.

†Number of seroconversion/number seronegatives observed X fraction of year observed X 100.

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TABLE 2—Participation in Examination Cycles 1–7 of SFMHS by HIV Seronegative Men According to Risk Status* on Entry

	Examination Cycles							
	1		2	3	4	5	6	7
Risk Status*	No.	%	%	%	%	%	%	%
0–1	68	100	84	79	84	79	81	74
2–9	141	100	80	70	74	83	74	78
10 or More	202	100	90	84	83	88	83	84
Total	411	100	86	78	80	85	80	80

*Numbers of male sex partners reported by HIV seronegative study subjects for two years before entering the SFMHS.

the six months before each examination cycle is shown in Fig. 1. Among HIV seropositive men, the prevalence of insertive anal intercourse with two or more partners in six months declined from about 40 per cent during January through June 1984 to about 15 per cent during the six months January through June 1985. The prevalence of this practice remained essentially unchanged during the succeeding 18 months, dropping to about 5 per cent during the first six months of 1987. The overall decrease in prevalence during the 36 months of follow-up amounted to about 87 per cent.

Among HIV seronegative men, the prevalence of receptive anal intercourse with two or more partners in six months declined from about 15 per cent during the six months January through June 1984 to about 5 per cent during the last six months of 1984. Prevalence of this practice has remained low since then, with the last reported prevalence for the period January through June 1987 at about 3 per cent. The overall decrease in the prevalence of this risk factor during the period of observation has been 80 per cent.

Discussion

Monitoring HIV seroconversion rates in representative samples of defined populations is an essential element in understanding the extent of epidemic spread of infection, evaluating the effectiveness of control efforts, and providing data for predicting the future extension of the epidemic, including incidence of the acquired immunodeficiency syndrome (AIDS) in the sampled population. This is the second report of HIV seroconversion rates from such a sample. The findings, however, will be of interest not only to those responsible for controlling the San Francisco epidemic, but also for other communities where sexual transmission of HIV infection among homosexual/bisexual men is a major problem.

The data reported here indicate that the previously reported decline in the spread of infection, from an estimated annual rate of approximately 18 per cent per year between 1982 and 1984 to approximately 4 per cent per year during the first six months of 1986,[1] has continued. For the latest period of observation, July through December 1987, the annualized HIV seroconversion rate in the SFMHS cohort was 0.7 per

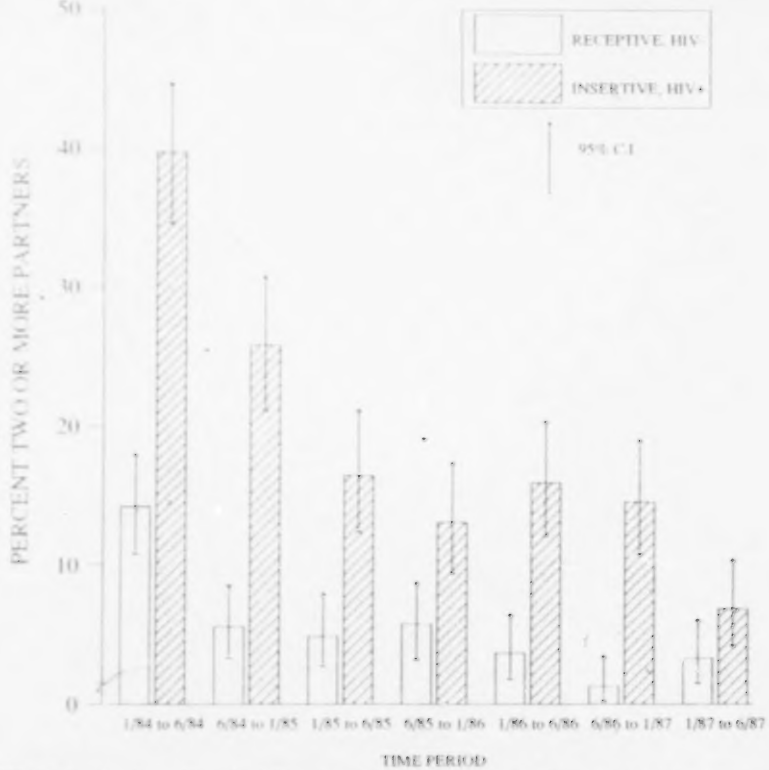

FIGURE 1—Prevalence of Sexual Behaviors Associated with Acquisition and Dissemination of HIV Infection among Homosexual Bisexual Men, SFMHS, January 1984–December 1987

cent. This low rate of infection was unlikely to be attributable to selection bias of study participants. A more likely explanation is the decline in the prevalence of the major modes of transmission, insertive anal intercourse among HIV seropositive homosexual/bisexual men, and receptive anal intercourse among HIV seronegative homosexual/bisexual men.[1] The decline in HIV infection rates in San Francisco has also been demonstrated in a cohort of homosexual/bisexual men originally recruited from sexually transmitted disease clinics for studies of hepatitis B infection and followed from 1978 to the present.[4]

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Next-of-Kin Attitudes Regarding Participation in an Epidemiologic Case-Control Study

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Abstract: We assessed next-of-kin's attitudes about participating in an epidemiologic case-control study of adult acute nonlymphocytic leukemia. Responses from a mailed questionnaire indicated that 95 per cent were glad they participated. While 74 per cent benefited, 18.5 per cent were bothered in some way. Results concerning the need to obtain physician permission before contacting next-of-kin were inconclusive; however, 8 of 10 females contacted by the case's physician considered consent necessary. (*Am J Public Health* 1988; 78:1474–1476.)

Introduction

Case-control studies have become an increasingly popular method in epidemiologic research. Depending upon the disease under study, some eligible cases may die before a personal interview can be conducted. In those instances, investigators may choose to ask the next-of-kin to provide information that would otherwise be obtained directly from the case.

While living subjects have been asked about their attitudes toward participating in epidemiologic case-control studies,[1,2] little empirical data is available to evaluate the opinions of interviewed next-of-kin. The present survey examines next-of-kin's attitudes about their participation in a population-based case-control study of adult acute nonlymphocytic leukemia (ANLL).

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Methods

The present survey was conducted as the final phase in an investigation of ANLL and residential exposures to power-frequency magnetic fields in King, Pierce, and Snohomish counties of Western Washington State. Eligible cases were identified through a population-based SEER (Surveillance, Epidemiology, and End Results) registry and included 164 individuals aged 20 through 79 who were diagnosed with ANLL from 1981 through 1984. Personal interviews took place between November 1984 and February 1986. Of the 139 cases who had died prior to being contacted by study investigators, interviews were successfully completed with 94 (67.6 per cent) of their next-of-kin (Table 1).

Each respondent could be asked to participate in two different parts of the study. First, a one hour personal interview was conducted in the next-of-kin's home. If the case had lived at this residence one or more years before her/his diagnosis date, the interviewer also took magnetic field measurements in the kitchen, bedroom, and family gathering room. Second, scientific instruments were installed in a limited sample of homes for the purpose of measuring magnetic fields over a 24-hour period of time.

TABLE 1—Next-of-Kin Response Rates for Original Interview and Mailed Questionnaire

Outcome	Original Interview		Mailed Questionnaire	
	No.	(%)	No.	(%)
Identified as eligible	139	(100.0)	94	(100.0)
Physician refusal	29	(20.9)	NA	
Next-of-kin refusal	8	(5.8)	12	(12.8)
Moved/unable to locate	6	(4.3)	0	(0.0)
Other reason	2	(1.4)	1	(1.1)
Total completed interviews/ questionnaires	94	(67.6)	81	(86.2)

NA, not applicable

One month after the completion of the personal interview, a two-page self-administered questionnaire was mailed to each next-of-kin in which they were asked to assess their feelings about participating in the study. Simple descriptive statistics have been used to present the results of that survey.

Results

Of the 94 next-of-kin interviewed in the original study, 81 (48 spouses and 33 other relatives) completed the mailed questionnaire (Table 1). Seventy-seven (95 per cent) of the respondents indicated that they were glad to have participated in the study. In answering the question about why they agreed to be interviewed, the vast majority expressed the hope that their participation might help others who have cancer (Table 2).

A high percentage of the next-of-kin said they benefited from their participation compared to a small percentage who were bothered (Table 2). Of those who benefited from their participation, 15.0 per cent were nonetheless bothered by some aspect of the study. In contrast, 33 per cent who received no benefits indicated that they were bothered.

When next-of-kin were asked whether they felt it was necessary for the investigators to obtain consent from the case's physician prior to contacting them, a slightly higher proportion thought physician permission was not necessary (35 versus 28, with 18 professing not to know). The attitudes of next-of-kin, presented in Table 2, varied little by sex, relationship to the case, combined family income, number of years the next-of-kin knew the case, number of months between death and the interview, or interview type (with or without magnetic field measurements). However, females, spouses, and respondents whose interviews included magnetic field measurements were more likely to feel that physician permission was necessary.

Sex also seemed to influence who the physician told about the study; the 10 next-of-kin contacted by the case's physician were all female (nine spouses and one daughter). While eight of these 10 females felt that physician consent was necessary, only five of 25 males and 15 of 44 females who were not contacted by a physician agreed.

TABLE 2—Perceived Feelings about Participation in Study

	No.	(%)
Why participated in study		
It might help other people who have cancer	75	(92.6)
I might be able to get some information	23	(28.4)
It might be interesting	19	(23.5)
I might find comfort in discussing my feelings with someone	12	(14.8)
My family or friends thought it was a good idea	7	(8.6)
Other reasons participated	8	(9.9)
Reasons benefited		
No benefits	15	(18.5)
Unknown benefits	6	(7.4)
Possible improvement in future medical care for others	48	(59.3)
Satisfaction of helping others	40	(49.4)
Educational experience	17	(21.0)
Interesting experience	14	(17.3)
Comfort in discussing my feelings with someone	10	(12.3)
Other benefits	2	(2.5)
Reasons bothered		
Not bothered	64	(79.0)
Unknown if bothered	2	(2.5)
I had inadequate information to answer the questions	5	(6.2)
Some questions were tedious or boring	3	(3.7)
It brought back painful, depressing memories	3	(3.7)
Some questions were too personal	2	(2.5)
There were no questions about leukemia	2	(2.5)
The interview took too long	1	(1.2)
Other reasons bothered	4	(4.9)

Discussion

The results from this study demonstrate that very few next-of-kin were upset by their participation. In fact, the vast majority benefited from their involvement and were glad to have been contacted. This corroborates the reports concerning the opinions of living cases and controls,[1–4] and extends those results to the experience of next-of-kin as research participants.

Results regarding the need to obtain physician consent are inconclusive. In most epidemiologic studies, the next-of-kin are not approached until some months after the case's death. Since the physician may not see the next-of-kin after the case dies, she/he may not be able to make an informed decision concerning the advisability of next-of-kin participation in a research study. Institutional review board policy that requires physician consent to contact next-of-kin may lead to what Lebacqz and Levine[5] and Rothman[6] describe as overprotection of the individual and abrogation of her/his privilege to contribute to expanding medical knowledge.

Boring, *et al*, reported that female cases whose doctors had contacted them were more likely to believe that physician consent was necessary;[2] the present study found similar results for female next-of-kin. Moreover, spouses and respondents whose interviews included magnetic field measurements were also more likely to feel physician permission was necessary. The nature of the next-of-kin's relationship to the deceased case and the type of interview may influence how the respondent feels about her/his participation.

These data should be interpreted with caution. First, only 58.3 per cent of the original next-of-kin pool responded to the mailed questionnaire. In addition to a limited sample size, it is possible that respondents with negative attitudes were excluded by the case's physician or their own refusal to participate. Second, the next-of-kin's feelings may vary based on the clinical course of the disease. This study relates to ANLL only. Additional studies may be useful in evaluating the effects of various diseases on the next-of-kin's feelings. Pre- and post-interview attitude surveys might also be of interest. Such information could provide valuable insights for institutional review boards as they assess the potential impact of proposed research on study participants.

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Inadequate Utilization of Routine Electronic RBC Counts to Identify Beta Thalassemia Carriers

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Abstract: We investigated physician awareness of the diagnosis of beta thalassemia minor as suggested by RBC indices obtained from routine electronic counts; and, the knowledge of the carriers of their genetic trait. Out of 17,000 counts, 324 were compatible with the diagnosis of beta thalassemia minor, but, only 175 (54 per cent) were identified by physicians as possibly thalassemic and in 47 of these (27 per cent) was the diagnosis verified. Twenty-four of 39 interviewed patients in whom the diagnosis of beta thalassemia minor was established knew about their carrier state. (*Am J Public Health* 1988; 78:1476–1477.)

Introduction

Thalassemia is the most prevalent hereditary hemoglobinopathy afflicting millions of individuals worldwide. It is a genetic disorder affecting synthesis of the globin moiety of the hemoglobin (Hb) molecule. The beta thalassemias are prevalent in a broad belt extending from the Mediterranean basin to Southeast Asia. In Israel it is the most common hemoglobinopathy with an absolute gene frequency of 0.1.[1] Beta thalassemia major denotes the homozygous, severe, transfusion dependent form of the disease (Cooley's anemia) of which 100,000 children die annually worldwide.[2] Beta thalassemia minor, on the other hand, is associated with mild or absent anemia but with reduced mean corpuscular volume (MCV) and mean corpuscular hemoglobin (MCH) and elevated HbA_2. The majority of thalassemia minor individuals, who are heterozygous for the beta thalassemia gene, are asymptomatic and unless diagnosed and informed may be unaware of their carrier state.

A basic tenet in the overall management of the beta thalassemia, as for any autosomal recessive disorder, is the prevention of its homozygous form through the detection and education of the heterozygous carriers. One way of achieving this goal is to screen populations at risk and instruct the identified heterozygous carriers about the genetic implications of marrying another carrier.[3] Of all screening tests so far advocated for beta thalassemia minor, electronic counts of red blood cells (RBC) indices is the best.[4–7] Electronic RBC counts is a rapid, automated, inexpensive and widely used test. The most striking and consistent finding in heterozygous beta thalassemia is the combination of a relatively high RBC count with a low MCV and MCH along with normal Hb and hematocrit values.[8–13] In fact, the MCV alone is a sensitive parameter, identifying a high number of thalassemia carriers both in adults and children (99.3 and 98.1 per cent, respectively).[14] In an adult Mediterranean population, the percentage of false negatives was as low as 0.7 per cent.[14] Since electronic cell counts are now obtained routinely worldwide, entire populations are screened de facto for thalassemia at no additional cost to the medical system. The present study was designed to assess the awareness of physicians to the possible diagnosis of beta thalassemia minor in adults as suggested by typical RBC indices, and the knowledge of diagnosed carriers about their status and its significance.

Methods

Medical records of the first 17,000 adults whose charts contained complete electronic blood counts were reviewed: 12,920 counts of patients hospitalized in Hadassah University Hospital, Jerusalem, and 4,080 gathered from associated community clinics. The counts were examined for compatibility with the diagnosis of thalassemia trait; criteria for compatibility were an RBC count of more than $6 \times 10^6/mm^3$ ($6 \times 10^{12}/L$), MCV less than 72 femtoliter and MCH less than 25 picogram, with normal Hb and Hct values. When RBC indices compatible with the diagnosis of thalassemia minor were found, the charts were searched for any note relating to this abnormality. When no mention was noted, the abnormal results were deemed ignored. When they were noted and thalassemia minor suspected, we looked for verification of the diagnosis: demonstration of elevated HbA_2 and/or HbF concomitant with normal or high serum iron.

Patients in whom the diagnosis of beta thalassemia minor was established were interviewed individually in order to determine their knowledge of the disease. The standardized interview included questions regarding awareness of the trait, its genetic implications, prenatal diagnosis, and symptomatology of the homozygous form.

Results

In 324 of the 17,000 (1.9 per cent) records reviewed, the electronic RBCs were compatible with the diagnosis of beta thalassemia minor. However, in only 54 per cent (175/324) was any mention of the abnormality noted in the chart even though all 324 individuals belonged to ethnic groups at risk for beta thalassemia (Iraqi or Kurdish extraction). The proportion of recognized abnormal indices was higher (62 per cent) in records of hospitalized patients compared with that of community-based charts (38 per cent).

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valid *legal* reasons for physician-initiated blood alcohol concentrations. When blood alcohol concentrations are used to provide realistic individualized patient guidelines regarding drunken driving, legal issues become moot: the patient's health becomes the only motivating factor, and the use of blood alcohol concentrations becomes equivalent to any other laboratory test obtained for diagnosis and treatment.

It is possible, given a blood concentration, to predict on average how long it will take for the level to decrease to a specified level.* This information can be used to provide impaired patients with a written "driving prescription" indicating the minimum time during which driving should be avoided. Since patients will vary in their metabolism of alcohol, the "driving prescription" should be conservatively based on time to a blood alcohol level of zero. The written prescription serves several useful purposes: it encourages physicians to provide realistic advice tailored to the individual patient; serves as a reminder for the sobering patient should they forget verbal instructions; should be medically acceptable; and demonstrates physician concern with the health consequences of driving while impaired.

Any diagnostic test should be obtained only for medical reasons that might lead to some benefit for the patient. This implies that action will be taken once the results of the test are known. More liberal use of blood alcohol concentrations might lead to advice for impaired patients, who are not obviously intoxicated, to refrain from driving. Such advice, if translated to fewer traffic crashes, would improve patient care. Since driving while impaired creates tremendous health risks, an explicit "driving prescription" may be an important medical intervention that could also promote highway safety.

*The average time to a specified alcohol level, along with other relevant parameters regarding alcohol drinking, can be predicted using the "Alco-Calculator (Alcohol Research Documentation, Inc., Rutgers University Center of Alcohol Studies, New Brunswick, NJ). The calculation assumes drinking has stopped sometime prior to the obtained blood alcohol concentration and is an estimation based on the population average.[10]

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Prescribing of Noncontraceptive Estrogens and Progestins in the United States, 1974–86

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Abstract: This paper describes changes in the prescribing of noncontraceptive estrogens and progestins, using data from pharmaceutical marketing surveys. The number of estrogen prescriptions decreased from 1975 to 1980, and then increased through 1986. Progestin use has increased since 1982; concomitant use of estrogens and progestins increased over time and was common in 1986. The trends suggest that the use of estrogens, particularly the combined use of estrogens and progestins, will continue to increase. (*Am J Public Health* 1988; 78:1478–1481.)

Introduction

Both positive and negative health effects of estrogen use have been reported.[1–5] Nevertheless, it is important to be aware of changes over time in the use of estrogens and of the characteristics of women exposed in order to estimate the public health importance of any effects. The notable changes in the quantity of prescribing of noncontraceptive estrogens in the United States from 1966 to 1983 has recently been described.[6] This paper describes the changes in the patterns of and indications for prescribing of noncontraceptive estrogens from 1974 to 1986.

Methods

Data were derived from two pharmaceutical marketing research data bases purchased by the Food and Drug Administration (FDA) from IMS America, Ltd, Ambler, PA: the National Prescription Audit (NPA) and the National Disease and Therapeutic Index (NDTI).[6,7] The NPA gives national estimates of prescription volume based on prescriptions dispensed by a panel of retail pharmacies. The NPA data presented are for all dosage forms of estrogen and include estrogen/androgen combination products.

The NDTI provides descriptive information on disease patterns and treatment in office-based practice in the US. A

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panel of over 2,000 physicians report on each patient seen or contacted (patient *visit*) in any way (office, hospital, telephone) during a 48-hour period each quarter. There are no published reports on the quality of the NDTI data, but we have assumed the possible biases to be similar over time, and the data valid for secular comparisons. Drugs reported during a patient visit ("mentions") may have been given as a formal prescription, hospital order, directly dispensed or administered, or just recommended. Concomitant use of drugs means drugs prescribed for the same diagnosis, and does not include drugs ordered for the same patient for different diagnoses. When regional variation was studied, the total number of drug mentions regardless of age was divided by the estimated number of women aged 45–64 years in the region.[8,9,10]

Results

The previously reported increase in estrogen prescription volume from 1980 to 1983 has continued through 1986, (Figure 1). In 1983, 16.6 million prescriptions for estrogen-containing products were dispensed, and in 1986 20.3 million: 37 per cent new and 63 per cent refills. Oral estrogens accounted for 88 per cent of estrogen prescriptions in 1986, with Premarin® still the most commonly used preparation. Vaginal estrogen products show little variation over time (2 million in 1986), and androgen/estrogen preparations are down from 7 per cent in 1974 to 1 per cent in 1986.

According to the NDTI data, the use of injectable estrogen has decreased from 5.0 million doses in 1974 to 1.8 million doses in 1986 while the use of oral estrogens increased 117 per cent between 1979 and 1986. Because each injection requires a patient visit, they get disproportionately more emphasis in the NDTI than oral estrogens. To minimize confounding that might be caused by the decreased use of injectable estrogens over time, further NDTI data on time trends will mainly concern oral estrogens.

The rate of prescribing of oral estrogens varied notably by region in the US, being lowest in the East and highest in the West (Table 1). Differences between the areas increased from 1974 to 1985, but the rankings remained the same. Similar differences were found for all estrogens. The share of oral estrogens prescribed by gynecologists rose while that prescribed by general/family practitioners declined from 1974 through 1986.

Oral estrogens were used mainly in women (92 per cent in 1986), aged 40–59 years (Table 2). In different years from

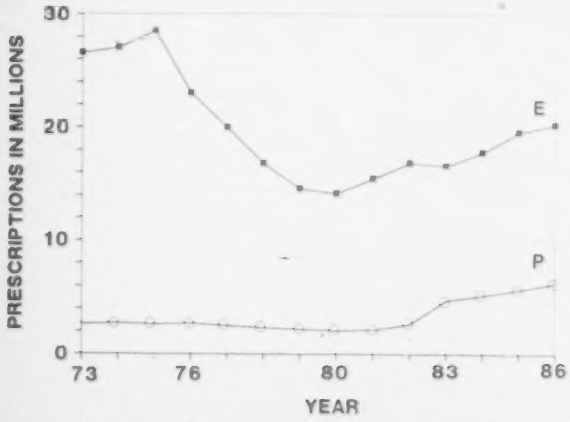

FIGURE 1—Numbers of Dispensed Prescriptions Containing Estrogens (E) and Progestogens (P) in 1973–86, in Millions (the NPA data).

TABLE 1—Rates of Total Oral Estrogen Mentions per 100 Women Ages 45–64 Years by Region, United States, 1974–85

Region	1974	1976	1979	1982	1985
	Rates/100 Women				
East	30	23	11	11	15
Midwest	38	29	14	25	27
South	34	19	15	17	26
West	45	36	18	31	43

SOURCE: National Disease and Therapeutic Index (NDTI), IMS America and US Department of Commerce.

66 to 82 per cent of oral estrogens were prescribed for problems associated with menopause or post-menopausal changes. The other important indication was cancer.

Other drugs have been frequently prescribed along with an estrogen (Table 3). Concomitant use of progestins increased substantially during the 1980s.

The NPA shows that the total number of progestin (P) prescriptions rapidly increased between 1984 and 1986 (Figure 1). In 1979, oral progestins were given mainly to women under age 40 (71 per cent of use), they were used alone (79 per cent), and for first-time therapy (81 per cent); menopausal problems represented only 18 per cent of the diagnoses. In 1986, oral progestins were given to women 40 years and older (68 per cent), only 37 per cent were used alone, 52 per cent were new therapy, and menopausal indications represented 59 per cent of diagnoses.

"Menopausal symptoms" was the most common diagnosis for which estrogens were prescribed (from 32 to 46 per cent of all estrogen use each year), as shown in Table 4. A drug was prescribed in most of the visits for "menopausal symptoms", and the use of oral estrogens and progestins in the treatment of "menopausal symptoms" has increased.

The number of visits with the diagnosis "osteoporosis" was 1.4 million in 1974 and 1.8 in 1986. The majority of visits (from 86 to 96 per cent) were made by women, and in 1974–82 most of them were age 60 or older. Since 1982, the proportion of visits by women ages 40–59 years have increased (up 17 per cent in 1986). The importance of non-combination estrogens in therapy increased from 1974 to 1986, and in 1986 over one-fourth of the patients received oral estrogens. However, calcium supplements were the most commonly prescribed therapy throughout the 1980s.

Discussion

The increase in estrogen prescriptions which began in 1980 continued through 1986, but the highest level was still

TABLE 2—Per Cent Distribution of Oral Estrogen Mentions by Patient Age, 1974–86

Age Groups[a]	1974	1976	1979	1982	1985	1986
0–19 (years)	1	2	2	—	1	1
20–39	17	17	22	21	11	13
40–59	67	64	57	60	63	62
60–64	6	8	6	9	9	10
65+	8	9	13	10	16	15
TOTAL	100	100	100	100	100	100

SOURCE: National Disease and Therapeutic Index, IMS America.
a) Females only

TABLE 3—Per Cent Distribution of Mentions of the Concomitant Use of Oral Estrogens with Selected Other Drugs, 1974–86

Estrogen Use	1974	1976	1979	1982	1985	1986
Used alone	65	67	71	63	53	52
With oral progestogens	2	5	4	9	23	28
With calcium	—	—	2	3	6	5
With benzodiazepines	5	3	3	2	2	2

SOURCE: National Disease and Therapeutic Index, IMS America.

somewhat lower than in the peak year of 1975. The increase in the prescribing of progestins started somewhat later (after 1982) and coincided with a trend toward the use of an oral progestin along with an oral estrogen. Injectable forms of estrogens and estrogen/androgen combinations have lost their popularity with time. These changes may be due to changes in prescribing habits or in the specialty of prescribing physicians.

As the overall prescribing went down, prescriptions for menopausal symptoms were most affected, prescriptions for surgical menopause and cancer less affected. This is also reflected in the sex and age distribution of patients receiving estrogens.

When the number of NDTI drug mentions for estrogens dropped by 59 per cent from 1974 to 1979, there was only a small drop (16 per cent) in the proportion of women with "menopausal symptoms" who were treated with oral estrogens. In contrast there was a large drop (64 per cent) in the number of visits in which this diagnosis was given. There are two possible explanations for this finding: it may be that treatment determined recording of diagnosis, and visits not having any prescriptions were underreported in the NDTI data base; or that women changed their help-seeking behavior, and the changes in the number of visits were real.

Osteoporosis as a distinctive diagnosis in estrogen therapy was rare although it became somewhat more common over time. However, the classification of reasons for visits in NDTI emphasizing the underlying problem rather than the aim of the treatment may underestimate the prevention of osteoporosis as a reason for estrogen therapy.

Consistent with earlier studies,[11,12] this survey showed that if a woman consults a doctor because of menopausal problems, she is very likely to get a prescription for estrogens. Given the purported benefits of estrogens and combined estrogen-progestin treatment, including claims that they are protective against cardiovascular diseases, and current concerns about osteoporosis, it appears that increased prescribing of estrogens will continue for some time.

TABLE 4—Treatment of Women with the Diagnosis of "Menopausal Symptoms", 1974–86

Patient Visits	1974	1976	1979	1982	1985	1986
Estimated visits (in millions)	9.4	7.2	3.4	4.1	5.7	6.2
	Per Cent of Visits					
Visits with drugs	89	83	78	81	80	81
Type of drug[a]						
Oral estrogens	37	33	31	39	45	52
Injected estrogens	36	31	27	25	18	14
Injected androgen-estrogens	7	7	5	4	3	3
Oral progestogens	1	2	2	6	16	20
Benzodiazepines	6	6	4	4	2	2
Vitamin B12	4	5	3	6	2	2

Source: National Disease and Therapeutic Index, IMS America.
a) Only drugs prescribed in at lesat 5% of visits in any year are included.

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Decreased Incidence of Hip Fracture in Hispanics, Asians, and Blacks: California Hospital Discharge Data

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Abstract: We examined the incidence of hip fracture in Non-Hispanic White, Hispanic, Black, and Asian Americans for the years 1983 and 1984 using a data base which contains a summary of all hospitalizations for the state of California. We found a consistently lower risk for hip fracture after age 60 in Hispanic, Black, and Asian American females than in White females who were not Hispanic. Overall age-adjusted hip fracture rates in Hispanic, Black, and Asian females were 49.7, 57.3, and 85.4, respectively, and 140.7/100,000 in White females who were not Hispanic. These differences were not found in males, although Whites (not Hispanic) had the highest incidence of hip fractures among males. (*Am J Public Health* 1988; 78:1482–1483.)

Introduction

Hip fracture, a major outcome of osteoporosis, is a significant cause of disability and mortality in the United States.[1]

There is relatively little published data on the incidence of hip fractures in ethnic groups. Previous epidemiologic studies have shown White females to have a twofold increased risk for hip fracture as compared to Black females[2] or Mexican Americans.[3]

Using a statewide hospital discharge data base for the State of California in the years 1983 and 1984, we estimate the incidence of hip fractures in four ethnic groups: Hispanics, Blacks, Asian Americans, and Whites who are not Hispanic.

Methods

The California Discharge Data Program under the administration of the California Health Facilities Commission maintains records of all hospital inpatient discharges from acute care hospitals in the State.[4] The data base was complete for the years 1983 and 1984. Race was categorized as White, Black, Hispanic, native American/Eskimo, Asian, other, and unknown. There were 3,551,874 hospitalizations in 1983 and 3,538,094 in 1984. Hip fractures were identified by the principal diagnosis of either capsular or intertrochanteric fracture, ICDA-9 codes 820.0 to 820.9.[5]

Age-specific and age-adjusted hip fracture incidence rates were calculated specific for sex and ethnicity for 1983 and 1984, and then averaged. The 1983 and 1984 population in each ethnic group was estimated from data derived from the US Census for 1980[6] and the Center for Continuing Study of the California Economy for 1985.[7] Age-adjusted incidence rates were calculated using the total 1980 California population as the standard, with the 10-year intervals of age.

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Results

In 1983, there were 17,556 hospital discharges for hip fractures in females, of which 16,898 occurred in White (not Hispanic), Hispanic, Black, and Asian females (143 were classified as American/Indian/other, and 515 as race unknown). Comparable figures in 1984 were 15,626 and 15,289 (134 American Indian/other, and 203 race unknown).

Age-adjusted and age-specific incidence rates of hip fracture within each ethnic group are shown in Table 1. White (not Hispanic) females had the highest age-adjusted rate at 140.7, while Hispanics, Blacks, and Asians were lower, at 49.7, 57.3, and 85.4, respectively. The incidence rate of hip fracture increased exponentially with age in all four ethnic groups.

In 1983, there were 5,914 hospital discharges for hip fractures in males. Of these, 5,729 occurred in White (not Hispanic), Hispanic, Black, and Asian males (40 were classified as American Indian/other, and 145 as race unknown). Comparable figures in 1984 were 5,292 and 5,185 (49 classified as American Indian/other, and 58 as race unknown).

Male hip fracture incidence rates within each ethnic group are shown in Table 2. White males (not Hispanic), like their female counterparts, have the highest age-adjusted incidence rates at 48.9/100,000, while Hispanics, Blacks, and Asians are lower at 22.2, 38.7, and 26.3, respectively. The rate of increase in incidence of hip fracture in males is initially less than that of females but beginning in the sixth decade there is a discontinuity with a rapid exponential increase thereafter. The hip fracture rate of Black males is similar to White males until the seventh decade. The hip fracture rate of Hispanic and Asian males remains consistently lower than White males in all decades.

The age-specific incidence rates for men are greater than those for women in the first five decades of life, with female-to-male ratios of 0.42 to 0.54. After age 60, the relationship reverses in all groups except Blacks, with female-to-male ratios varying from 11.7 to 2.3. By age 70, sex ratios are higher for females in all ethnic groups, ranging from 1.4 to 2.2. By the oldest ages, the incidence rate for men has increased sharply but still trails the incidence rate for women, with female-to-male ratios of 1.2 to 2.6.

Sex ratios based on age-adjusted overall rates of hip

TABLE 1—Hip Fracture Incidence Rates (hospital discharges/100,000 person years) among Females by Ethnic Group and Age (1983 and 1984 averaged)

	Females			
Age (years)	White (not Hispanic)	Hispanic	Black	Asian
<50	6.2	2.6	5.3	2.7
50–59	65.3	15.7	35.2	17.4
60–69	212.9	59.6	80.7	91.8
70–79	726.0	246.0	267.8	321.6
80+	2502.3	958.7	988.8	1912.7
TOTAL	140.7	49.7	57.3	85.4

TABLE 2—Hip Fracture Incidence Rates (hospital discharges/100,000 person years) among Males by Ethnic Group and Age (1983 and 1984 averaged)

	Males			
Age (years)	White (not Hispanic)	Hispanic	Black	Asian
<50	11.5	5.6	12.4	6.3
50–59	37.3	15.0	46.4	16.4
60–69	91.5	34.3	83.6	49.1
70–79	334.1	145.9	189.2	155.0
80+	1209.6	595.1	816.4	738.7
TOTAL	48.9	22.2	38.7	26.3

fracture were highest in Non-Hispanic Whites and Asians (2.9 and 3.3, respectively) and lower in Hispanics and Blacks (2.2 and 1.5, respectively).

Discussion

Our estimated incidence rates were subject to errors of ascertainment of the number of cases (the numerator) as well as the total susceptible population (the denominator). The number of cases may have been under- or overestimated by misclassification of disease or ethnicity. Although general validity studies have been done by the State of California, the investigators were unable to validate the specific ICDA codes used in this study due to the confidential nature of the data base.

Significant errors in ascertainment of ethnicity of individual cases may occur. In 1983, 3 per cent of hip fracture cases had unknown ethnicity (1.3 per cent in 1984).[4]

The number of cases would be underestimated if patients were treated outside of a hospital, but such events are rare since hip fracture usually necessitates hospitalization.

The number of cases may have been overestimated due to readmissions or admissions due to a second fracture in the same year. Readmission is possible but the ICDA code for principal diagnosis should specify a complication of the fracture or operative intervention. Previous studies have found the incidence rate of second hip fracture to be 9 per cent in a 10-year period.[8] This suggests that refracture would increase the incidence rate at less than 1 per cent per year in patients with prior fracture.

Some errors in the estimation of the total susceptible populations probably occurred. In this study, the likely effect of underestimation of the denominators would be to decrease differences between Whites and other ethnic groups. Thus, our estimates of ethnic differences may be underestimated.

Risk factors for hip fracture include dietary intake,[9] reproductive history,[9] anthropometric data,[9] and bone-mineral content.[9,10] We observed striking differences between ethnic groups in their incidence rates of hip fracture. However, some of these known risk factors cannot easily explain these ethnic differences. Hispanics have a lower dietary intake of calcium than Whites[11] but a lower rate of hip fracture. Asians have lower bone mineral content than Whites,[12] yet Asians in our study had a decreased risk. The time of menopause is similar in all ethnic groups.[13]

We conclude that both male and female Hispanics, Blacks, and Asians have lower risks of hip fracture than do Whites who are not Hispanic. Further studies are needed to define differences in risk factors between these ethnic groups. Information concerning differences in hip fracture rate and risk factors may be helpful in designing prevention strategies for different ethnic groups.

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The Stability of Health Status in Rheumatoid Arthritis: A Five-Year Study of Patients with Established Disease

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Abstract: We employed a health status measure to describe the outcomes of rheumatoid arthritis patients over five years. Of the 410 rheumatoid arthritis patients who were originally administered the Arthritis Impact Measurement Scales (AIMS), 299 completed a follow-up five years later. Data were analyzed using nine health status scales, three components of health status, and an overall arthritis impact item. Results for survivors indicated that there were no clinically important deteriorations in any of these measures. In fact, small improvements on most measures were found. The health status changes were similar for patients originally in a clinical trial and for those receiving routine specialty care. Age was found to positively relate to improvements in psychological status and overall arthritis impact, but we were unable to demonstrate any consistent effects of sex, marital status, education, or disease duration. Our results contrast with other studies that have noted major declines over time in the health status of patients with rheumatoid arthritis. Furthermore, level of education was not a major determinant of morbidity in this group. These results suggest that health status in certain patients with rheumatoid arthritis is more stable than previously thought. This has implications for both clinical practice and clinical research in rheumatology. (*Am J Public Health* 1988; 78: 1484–1487.)

Introduction

Rheumatoid arthritis (RA) is a chronic disease that generally leads to structural damage and functional disability. In the past, the course of rheumatoid arthritis was monitored in practice and described in the literature most often in terms of standard clinical variables such as joint count, range of motion, and x-ray erosions. Studies using these measures generally painted a picture of steady and severe decline over time.[1,2] More recently, functional status measures have been employed to study the impact of rheumatoid arthritis. In general, such studies also have suggested that functional loss over time is substantial.[3,4] Age, sex, duration of illness, and level of education have been shown to be predictors of disability in these studies.

This study used the Arthritis Impact Measurement Scales (AIMS) to describe health status changes over five years in a group of subjects with established rheumatoid arthritis. The AIMS is a reliable and valid assessment tool that has previously been shown to be sensitive to changes in health status in RA.[5,6] Our results indicate that health status in rheumatoid arthritis may be quite stable over time. We also found that sex, age, marital status, education, and disease duration were relatively weak predictors of changes in health status.

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Methods

The AIMS is a patient self-administered questionnaire that includes 64 items of which 45 are specific health status questions.[7] These items are grouped into nine scales which address the following aspects of health status: *mobility* (ability to move around the community); *physical activity* (lower extremity functions); *dexterity* (upper extremity functions); *household activities* (basic household tasks); *activities of daily living* (basic self care activities); *social activity* (social interactions with friends and family); *anxiety*; *depression*; and *pain*.

Previous work has demonstrated that these nine scales can be summarized into three broad components of health status: *physical disability* (which includes mobility, physical activity, dexterity, household activities, and activities of daily living), *psychological status* (which includes depression and anxiety), and *pain*.[8] The AIMS instrument also contains a single visual analogue scale that assesses overall arthritis impact. This item measures the patient's overall assessment of the effect of arthritis on their health status. All of the AIMS scores are rescored to a 0 to 10 range where lower scores represent better health status.

The AIMS questionnaire was originally administered to 410 patients with classical or definite rheumatoid arthritis between 1978 and 1982. Patients were gathered from a variety of rheumatology practice settings in 10 different states. Two different patient populations were involved: 51 per cent were from clinical settings where they were receiving routine specialty care while 49 per cent were initially involved in a 21-week cooperative drug trial which compared placebo, oral gold, and injectable gold.[9] Since the study was not designed to assess the effect of treatment, there were no restrictions on therapy and no data were collected on which drugs patients had taken.

A second AIMS administration was conducted by mail from 1983 to 1986 such that there was approximately a five-year lag between the first and last administrations for all respondents. Rigorous follow-up procedures were used to improve response rates including the tracking of lost patients through friends, relatives and neighbors by the last known residence recorded for a patient. We also examined death records by locality for patients who did not respond to the mailings. Patients not responding to any of three mailings and not located in the death registry was considered lost to follow-up. Of the 410 rheumatoid arthritis patients who were initially administered the AIMS instrument 299 patients completed a follow-up questionnaire five years later. This represented 73 per cent of the group. Among the non-respondents, 37 (9 per cent) were known to be deceased and 74 (18 per cent) were lost to follow-up either because they had refused to participate (4 per cent) or had not responded to any of the mailings (14 per cent).

Analysis of the data began with an examination of the baseline comparability between the patients in the completed study group, those who were lost to follow-up, and those who died. We also compared the baseline scores of those who were initially in the clinical trial with those who were not. These comparisons included the demographic variables, the

three health status components, and the overall arthritis impact score. The differences between the patient's AIMS scores on the two administrations were calculated, with a minus indicating more disability or worsening on a particular scale or component and a plus indicating improvement. Changes in the AIMS scores had a potential range of minus 10 to plus 10. The mean change and 95 per cent confidence intervals for each of these health status scores were estimated.

Preliminary bivariate associations between change scores and selected disease and demographic characteristics were obtained. Those variables with a significant association at the .05 level were included as explanatory variables in a multivariate regression analysis of the changes in health status scores. Next, the multivariate analysis was adjusted for initial baseline measures of physical status, psychological status, and pain. A final adjustment in the multivariate analysis was the inclusion of a dummy variable to distinguish the general and clinic study populations.

Results

The average age of the 299 patients who completed the follow-up was 52; they were predominantly female (69 per cent), and White (83.5 per cent); 36.9 per cent of the group had more than a high school education and 14.5 per cent were in professional occupations. The average disease duration in the group at the second administration was 14 years. Patients who completed the questionnaire were significantly more likely to be White, married, better educated, and have a higher income than those who died or were lost to follow-up. The baseline health status of the three groups was comparable, with the exception of the physical function of the deceased group which was worse. The routine care and clinical trial subgroups were similar for sex, marital status, education, occupation, income, and disease duration. However, the routine care group was older and more often White.

Baseline health status scores and change scores at five years for those who completed the second AIMS questionnaire are shown in Table 1. Baseline scores ranged from 6.30 for physical activity to 0.95 for ADL. The overall arthritis impact score at baseline was 4.36. No substantial deterioration was found at five years in any of the health status measures. In fact, almost all of the scales and components showed a trend toward slight improvement. Health status change scores were comparable for the clinical trial and routine care segments of the sample when adjusted for their initial baseline values.

A series of multivariate regression analyses were performed to examine the effect of selected study variables on changes in the health status components and arthritis impact scores. Four separate regression equations were constructed, one for each of the health status components and one for the impact score, with each equation being adjusted for the initial score. The results shown in Table 2 indicate that age, duration of disease, and marital status were each associated with significant changes in two of the four health status scores. Education was associated with changes in pain, being widowed was associated with changes in physical function, while sex did not have an effect on any of the scores. These findings demonstrate that none of the demographic variables had a strong or consistent effect across the spectrum of health status scores. There were no significant differences between the results for subjects in the routine care and clinical trial groups.

Discussion

We found little if any deterioration over a five-year period in the health status of patients with established rheumatoid arthritis. In fact, nearly all of the health status measures in this study showed slight improvement, although these changes were not clinically important. This finding of relatively stable health status was true for a group of patients who were initially receiving routine subspecialty care as well as for a group that was initially involved in a drug trial.

These results confirm work by Sherrer and colleagues at Stanford who reported that health status in rheumatoid arthritis is relatively stable after the first few years of disease.[10] In that study, it was reported that disability progressed most rapidly during the early years of disease and then assumed a slow, more linear rate of progression in later years. Over an average follow-up period of 5.6 years in a group of patients with an average disease duration of 10 years at entry,[11] the percentage of their patients with no disability was slightly higher than the percentage with severe disability

TABLE 1—Mean Changes in Health Status Measures by Scale and by Component over Five Years in 299 Subjects with Rheumatoid Arthritis

Scale/Component	Initial Score†	Change Score††	95% Confidence Interval for Change Scores
Scale			
Mobility	1.93	0.07	(−.26, .40)
Physical activity	6.30	0.41	(.10, .72)
Dexterity	4.85	0.42	(−.01, .85)
Activities of daily living	0.95	0.02	(−.20, .24)
Household activities	1.27	−0.05	(−.24, .16)
Anxiety	3.95	0.19	(−.05, .43)
Depression	2.82	0.25	(.03, .47)
Pain	6.18	0.96	(.67, 1.25)
Social activity	4.57	0.36	(.11, .61)
Component			
Physical	3.05	0.18	(−.04, .40)
Psychological	3.39	0.23	(.01, .45)
Pain	6.18	0.96	(.67, 1.25)
Arthritis impact	4.36	0.42	(.09, .75)

†Initial mean score based on score of 0 (best health) to 10 (worse health).
††Mean change score using score of −10 (maximum worsening) to +10 (maximum improvement).

TABLE 2—Regression Analysis of Changes in Health Status on Selected Study Variables

Study Variable	Changes in Health Status†			
	Physical	Psychological	Pain	Arthritis Impact
Intercept	.31	−1.78**	−.97	−2.51*
Sex				
Female	.29	−.05	.2[illegible]	.06
Male	—	—	—	—
Age (per year increase)	−.004	.02*	−.002	−.03*
Years of Arthritis (per year increase)	−.03**	−.01	−.02	−.04*
Marital Status				
Married	—	—	—	—
Separated	−.83*	−.31	−1.22*	−1.44
Divorced	.03	.05	.36	−.09
Widowed	−.91*	−.28	.17	.12
Never married	−.76*	.01	.86	1.67*
Education				
Less than high school	−.47	−.37	−.90*	−.25
High school	−.13	.09	−.04	.82
More than high school	—	—	—	—
F Value	4.56***	4.95***	5.73***	2.99***
R - Square	.18	.19	.19	.16

†Beta Coefficients are indicated for each study variable in each regression equation. Each regression equation includes covariate adjustments for initial baseline measures of physical status, psychological status and pain.
*p < .05
**p < .01
***p < .001

(12 per cent versus 11 per cent). Taken together, the findings of that study and this one strongly suggest that health status in rheumatoid arthritis is relatively stable in patients who have had the disease for more than a few years.

Our findings did not indicate that disease duration or demographic factors were consistent determinants of changes in health status. Age, marital status, number of years of disease, and level of formal education were nearly all associated with changes in one or at most two of the four health status scores over the five-year period. However, none of these factors, including level of formal education, proved to be a consistently strong determinant of changes in health status.

In particular, education level was not strongly correlated with change in physical disability. This finding stands in marked contrast to the recently published findings of Pincus and colleagues.[12] It is important to point out, moreover, that demographic variables in general are not very strong predictors of health status. In our study, regression models using demographic and disease duration variables explained approximately 18 per cent of the variability in the four health status scores. In the Stanford study, duration, age, and sex explained less than 25 per cent of the variability.

Our results suggest that health status of patients with established rheumatoid arthritis is more stable than previously thought, particularly if one takes a broad view of health that includes psychological status and pain in addition to physical function. Clinicians who treat patients with rheumatoid arthritis need to consider the pattern of health status change in this disease when deciding on treatment. Aggressive drug therapy in the early stages of disease may be justified since much of the decline in health status appears to occur during this phase. Conversely, patients with established disease may not be good candidates for therapeutic change in the absence of major symptoms or obvious health status declines.

Researchers involved in clinical trials of rheumatoid arthritis also need to consider the implications of these findings. Patients with early disease may be the best candidates for therapeutic trials from both the clinical and statistical perspective since substantial change in outcome is more likely to take place in this group. In addition, sample designs for such trials should seek to gather subjects within a rather narrow range of disease duration so that the results of the trial may be more easily analyzed and interpreted. Finally, and most importantly, broad measures of health status should be included as major outcome measures in any clinical trial of rheumatoid arthritis.

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Selection of Elderly Controls Using Random Digit Dialing

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Abstract: The experience of identifying and soliciting elderly male controls, using a nonclustered random digit dialing procedure in a case-control study, is presented. For elderly controls (ages 65–84) 3.5 times more residential telephone numbers were required than for controls encompassing a much broader age range (ages 40–84). This is a function of the proportion of elderly in the population and their lower response to telephone identification. Elderly controls, age 70 and older, also had lower participation rates. (*Am J Public Health* 1988; 78:1487–1488.)

Introduction

The results from random digit dialing varies among studies for many reasons including the sampling frame used, type of location (urban vs rural), geographical area of the country, age and sex of respondents, method of converting initial telephone screening refusals, and number, placement and degree of sensitive questions to determine eligibility.[1–9] The number of telephone numbers required to obtain controls in pediatric case-control studies has been reported.[8,9] The elderly have the highest proportion of telephone numbers of any age group in the United States.[7] However, the effort required to screen for elderly male controls using random digit dialing compared to younger adult males has not been reported. In this paper we present our experience with a random digit dialing procedure to solicit adult males.

Methods

This was a case-control study to explore dietary factors of potential etiologic significance in the occurrence of pancreatic cancer. Cases were ascertained from death certificates because one year survival is less than 15 per cent.[10] Controls, defined as living White males, ages 40–84, who resided in the seven-county Minneapolis-St. Paul metropolitan area, were frequency matched to cases and ascertained by random digit dialing. The interview was conducted with the control's wife (or other relative) since a surrogate was used to obtain information for the cases.

Initially, cases were divided into five-year age groups so that a similar number of controls could be obtained thus ensuring reasonable comparability on age. An attempt was made to select controls from similar geographic locations within the study area by incorporating the three digit prefix of the cases' telephone numbers into the random digit dialing procedure. The remaining four digits were randomly generated. For each case an initial list of 25 telephone numbers was available. These numbers were called in order and resolved (ineligible, refusal, etc.) before the next number on the list was solicited. Each telephone number was called up to nine times and the calls were evenly distributed among mornings, afternoons, and evenings and weekdays and weekends. If there was no contact after nine calls, the interviewer proceeded to the next telephone number.

An additional variation on random digit dialing was to check each telephone number in a reverse street directory to determine if the number represented a residence or commercial establishment. If a residence, a name was obtained so that the caller could properly address the respondent rather than having to solicit this information as part of the call.

At the initial call information was requested as to whether any males ages 40–84 resided in the household. If there was, a letter was sent which requested their participation. If not, the next number was called until an eligible, willing participant was ascertained. Residences that refused to provide screening information and eligible controls who refused to participate were contacted up to two additional times with letters and follow-up telephone calls. Because of the greater percentage of younger (ages 40–64) men in the population, this group was completed sooner than the older age groups (ages 65–84). Therefore, the age criterion during the enrollment period changed from 40–84 to 65–84. Results from the random digit dialing are presented according to the age groups screened: the initially screened broader age group (40–84) and the narrower, more elderly age group (65–84).

Results

Altogether there were 19,640 telephone calls made to determine the status of 8,936 telephone numbers used in the study (Table 1). The percentage of all telephone numbers called that were residences was similar between the broader age group (52.2 per cent) and the narrower, more elderly age group (55.6 per cent). The mean number of residential telephone numbers required to ascertain a control was 10.3 for the broader age and 34.3 for the more elderly age group. There was a higher percentage of residences that refused to provide the screening information using the broader age screen (6.7 per cent) than the more elderly age screen (5.0 per cent). On average, it took 5.5 hours to identify an eligible control. The following are the percentages (in parentheses) of

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TABLE 1—Random Digit Dialing Results

Outcome	Age Group Screened		
	40–84	65–84	Total
Total number of telephone calls needed to determine outcome	9,230	10,460	19,690
Total individual telephone numbers dialed	4,173	4,763	8,936
Numbers dialed:			
Non-working numbers	1,474	1,512	2,986
Non-residential numbers	513	613	1,126
Residential numbers	2,186	2,638	4,824
Ineligible residences	1,828	2,429	4,257
Number of refusals and eligibles:			
Number of residences which initially refused to provide screening information	260 (11.9)*	282 (10.7)	542 (11.2)
Number of initially refused residences which never provided screening information	146 (6.7)	132 (5.0)	278 (5.8)
Number of eligible residences identified	212# (9.7)	77 (2.9)	289 (6.0)
Mean number of telephone numbers per identified eligible residence	19.7	61.9	30.9
Mean number of residential telephone numbers per identified eligible residence	10.3	34.3	16.7

*Percentage of residential numbers in parentheses.
#Fifty-seven per cent were less than 65 years of age.

eligible controls who agreed to participate in the study: less than 60 years of age (80.2), ages 60–64 (82.2), 65–69 (92.6), 70–74 (72.7), 75 and older (69.5).

Discussion

This study demonstrated the greater difficulty in ascertaining only elderly male controls using a non-cluster random digit dialing technique. The non-cluster random digit dialing method has been discussed as straightforward and unbiased[8] although it will contain more telephone numbers that belong to nonresidences than does Waksberg's two-stage cluster method.[9]

Approximately 3.5 times as many residences were contacted in order to identify an eligible elderly control compared to a control from the broader age group. Because of the study design, we were unable to analyze separately the elderly controls ascertained through the use of the broader age range criterion. In spite of this limitation, the present study does illustrate the greater difficulty in identifying and soliciting elderly (ages 65–84) controls using random digit dialing. This is probably a function of the lower proportion of elderly in the population and their lower response to telephone identification.[7] Unfortunately, random digit dialing does not allow for the characterization of those who refuse to provide screening information.

Conceivably, if a cluster random digit dialing procedure was used, it could have reduced the number of calls needed to find the controls. However, there would still have been many more telephone calls necessary to obtain elderly controls. Eligible elderly controls age 70 and older also had a lower study participation rate.

The experience from this study suggests that random digit dialing may not be the most efficient method to identify and solicit elderly controls. An alternative method for soliciting controls among the elderly is through the Health Care Financing Administration which has been used successfully in case-control studies.[12] However, to use these records, the study must be federally funded.

ACKNOWLEDGMENTS

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Visits to Physicians Before and After Exposure to Urea Formaldehyde Foam Insulation

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Abstract: The average number of visits to a physician made by a sample of 351 residents of homes insulated with urea formaldehyde foam insulation in Montreal in the one year period before exposure was 5.25, and in the year following 5.62, an increase of 7 per cent (odds ratio 1.07, 95%CI=1.00,1.15). The increase in visits in the post insulation year was limited to subjects who had the product installed in the winter (OR=1.48, 95%CI=1.18,1.85), and was not seen for study subjects who insulated their homes during other seasons of the year. (*Am J Public Health* 1988; 78:1489–1491.)

Introduction

The possibility that exposure to urea formaldehyde foam insulation (UFFI) might be detrimental to good health was raised after publication in late 1979, of preliminary results of a study reporting development of nasal carcinoma in rats.[1] By the summer of 1980 homeowners and the general public became concerned about the health effects of UFFI.[2] Studies of residents experiencing high frequencies of allergic reactions, difficulty breathing, nasal stuffiness, headaches, and running eyes were quickly reported in both the scientific[3–5] and lay press, and there was speculation about possible long-term effects: cancer and congenital abnormalities.

Canada and several states in the United States banned the use of UFFI in 1981. Estimates of the number of homes insulated with UFFI are 500,000 in the United States and 100,000 in Canada.

We assessed the health impact of UFFI installation by exploiting the historical records of a universal health insurance system several years after the date of installation of UFFI.

Methods

Study subjects were drawn from the lists of the applicants to a provincial program which offered financial assistance to remove UFFI from homes. Established in 1983, the program required that the "individual in the household most affected by UFFI" be examined by a physician and a report made. There were 2,393 applicants from Montreal. The last group of applicants (453) were not coded by the government agency and were excluded. From the remaining 1,940 applicants we drew a stratified random sample of 351 subjects by sex and 10-year age groups. All subjects gave consent that the information could be used for research purposes.

Demographic information, date of UFFI installation, and Medicare number were obtained from the records of the government program. A subsequent telephone interview established if exposure to UFFI had ended at any time by moving out of the home with UFFI, by removal of UFFI, or through death.

Medical care utilization data were provided by the Régie de l'Assurance-Maladie du Québec (Régie), a universal health insurance system in Québec. Data were supplied for all instances of care received from November 1976 (when this particular type of data storage began) to January 1984. The physician visit required for application to the government program was excluded.

Paired t-tests and logistic regression analyses were done.[6] If the number of visits during two periods are Poisson random variables, then conditioning on the total number of visits during the two periods, the number of visits during one period is a binomial random variable with probability equal to the ratio of its Poisson mean to the sum of the two Poisson means. Thus if the mean number of visits are the same in the period before and after insulation, the proportion during the second period will be 0.5 and the odds ratio associated with exposure will be 1.0. Parameters for seasons of installation were included in the model to check whether the UFFI effect varied according to season of installation of UFFI.

We also adjusted for an annual baseline increase in number of visits to physicians by the general population of Quebec of about 3 per cent over the period of the study.[7] We took this into account by setting the offset in the logistic regression equal to the logit of 1.03, or $\log_e(103/203)/(1-103/203)$.[6]

Results

Of the total sample of 351, two residents were excluded because they had not insulated their homes with UFFI, five subjects because of coding errors, and seven with an unknown or incorrect Medicare number. Of the remaining 337 residents, 323 (96 per cent) completed the study interview (93 per cent of total sample of 349). All subjects with any missing information were excluded, leaving 303 subjects in the analysis (87 per cent of the initial 349 subjects selected for the study).

July 3, 1979 was the mean date of UFFI installation with most installations occurring in the 1978-80 period. The mean end of exposure date was April 1983 with a range of 1980 to 1984, when the final interviews were completed. The mean post exposure period was three years, nine months. The mean delay from time of UFFI installation to application for the assistance program was two and a half years.

Table 1 shows pre and post exposure period physician visits for study subjects. The average number of visits in the one year period before exposure to UFFI was 5.25, and in the year following UFFI installation was 5.62. The differences are accounted for by winter installations rather than installations in other seasons. The increase is most evident in the first and second seasons following initial exposure. Neither

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TABLE 1—Mean Number of Visits and Mean Differences[a] (and standard deviation) for Frequency of Visits to a Physician in Various Seasons after Installation of Urea Formaldehyde Foam Insulation (UFFI)

Season since Exposure	Mean Number of Visits		Mean Differences[a] in Visits (and standard deviation)				
			All Seasons	Winter	Spring	Summer	Autumn
0	Pre-insulation	1.33	.01	0.26	−0.39	−0.03	0.30
	Post-insulation	1.34	(2.30)	(1.36)	(2.35)	(2.28)	(2.63)
1	Pre-insulation	1.20	0.23	0.54	0.29	0.13	0.19
	Post-insulation	1.43	(2.42)	(2.08)	(2.75)	(2.04)	(2.89)
2	Pre-insulation	1.45	0.03	0.71	0.02	0.11	−0.45
	Post-insulation	1.48	(2.84)	(1.84)	(2.63)	(3.11)	(2.85)
3	Pre-insulation	1.26	0.11	0.20	−0.02	−0.01	0.38
	Post-insulation	1.37	(3.19)	(1.59)	(2.55)	(2.08)	(5.28)
Year before		5.25	0.37	1.71	−0.10	0.20	0.42
Year after		5.62	(6.32)	(4.32)	(6.03)	(5.33)	(8.66)

a) Differences in comparison to the same person in the same season, before, and after UFFI exposure.

increases nor decreases in visits are evident for those who had UFFI installed at other times of the year.

This same pattern is seen in Table 2 where odds ratios are presented corresponding to different seasons of UFFI installation. The overall increase in visits after UFFI installation in the winter months is 48 per cent, corresponding to an odds ratio of 1.48 (95% confidence interval = 1.18,1.85). Inclusion of calendar year in the logistic regression model indicated no modification of the UFFI effect by calendar year.

After adjusting the model for the 3 per cent inflationary annual increase in number of visits by the general population, results showed the same pattern with all odds ratio estimates reduced by about 3 per cent. For example, odds ratios for winter were OR=1.43, spring OR=0.95, summer OR=1.02, and autumn OR=1.03.

Discussion

This study shows that utilization of physician services by residents of homes insulated with UFFI increased somewhat in the year following winter insulation (January, February, and March). UFFI installation at other times of the year did not show a similar pattern. While such an effect could be a sampling artefact, it is consistent with the findings that chemical exposure from UFFI may occur as a result of incomplete hardening of the insulation when it is installed in the colder months.[8] Also, chemicals may reach higher concentrations in the interior of homes when ventilation is minimized, as might be the case during the winter.[9–11]

Previous epidemiologic studies on the health effects of UFFI have been limited by the absence of comparison groups,[12,13] possible bias in reporting behavior due to use of self-reported symptoms as an outcome measure,[12–16] and the difficulty of attributing symptoms solely to one environmental exposure.[12–16] By using self-paired comparisons and an objective measure of health status, this study avoids some of these limitations.

Nevertheless, several potential biases may exist in the present study. The use of a sampling frame of homeowners who make use of a government assistance program is not representative. It is likely that self-selection of residents who perceived that they were suffering from exposure to UFFI led to an increased possibility of finding an apparent effect.

In addition, there may have been underreporting of physician visits. However, research on a similar system in Manitoba has shown that the fee-for-service scheme provides valid data on total physician-patient contacts.[17]

Finally, with the records provided, it was not possible to assess the diagnoses for subjects. Possibly only more "serious" health problems which would prompt a person to visit a physician were assessed in this study.

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TABLE 2—Odds Ratios[a] for Frequency of Visits to a Physician in Various Seasons after Exposure to UFFI

Season since Exposure	Season Installed				
	All Seasons	Winter	Spring	Summer	Autumn
0	1.01	1.35	0.76	0.97	1.17
1	1.19	1.68 (1.05, 2.68)	1.21	1.12	1.12
2	1.02	1.83 (1.17, 2.86)	1.01	1.09	0.78
3	1.09	1.17	0.99	0.99	1.23
Year after vs year before	1.07	1.48 (1.18, 1.85)	0.98	1.04	1.06

a) Odds ratios for frequency of visits after exposure in comparison to the same person, same season in the previous year, with the null hypothesis that the frequency of visits was equal for each season before and after exposure.
Note: 95% CI in parentheses.

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Wife's Level of Education and Husband's Risk of Primary Cardiac Arrest

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Abstract: Data from a case-control study were analyzed to examine wife's education and risk of primary cardiac arrest in the husband. Men whose wives had more than 12 years of education had 80 per cent the rate of men with less educated wives (odds ratio = 0.8, 95% confidence interval = 0.5, 1.3), after adjustment for risk factors. There was no evidence of a status incongruity effect. These data are inconsistent with reports of positive associations between wife's education and coronary heart disease. (*Am J Public Health* 1988; 78:1491–1493.)

Introduction

Data from the Framingham Study revealed that the risk of coronary heart disease among married men was positively associated with their wives' education.[1] Reports from other populations suggest that this association may apply especially to men who have low education or who demonstrate Type A behavior.[2,3] During the past 20 years, investigations of direct and indirect socioeconomic measures in the United States and Europe typically have found an inverse association with coronary heart disease,[4–9] so a positive relationship with wife's education is a notable exception and may reflect the impact of stress-inducing changes in women's social roles and marital expectations on the home environment.[1,2]

To provide further information on wife's education as a risk factor, we analyzed data from a community-based, case-control study of one major form of coronary heart disease, primary cardiac arrest (PCA). Because this study was restricted to married individuals, it offered the opportunity to examine associations involving wife's education, separately and in combination with husband's education.

Methods

The study design and methods have been described in detail elsewhere.[10,11] PCA was defined as a sudden, pulseless condition without evidence of non-cardiac cause.[12] All cases (fatal and non-fatal) of out-of-hospital PCA in King County, Washington between December 1979 and January 1981 were identified by trained paramedics. The study was restricted to 25-75 year old married residents of King County who had no history of clinically recognized heart disease or activity-limiting co-morbidity (e.g., chronic lung disease, musculoskeletal impairment). For all study subjects, information was obtained in home interviews with spouses. After each case was identified, a control subject meeting the same criteria for eligibility and matched for age (± 7 years), sex, and urban or suburban residence was chosen by random digit dialing.[13] Of eligible subjects identified, 89 per cent of case spouses and 85 per cent of control spouses agreed to participate, producing 163 case-control pairs. Eighty-two per cent of the case-control pairs were male, slightly over half were urban residents, and the average age of cases and controls was 58 and 57 years, respectively.

This analysis was restricted to the 133 male case-control pairs. Education of the men and their wives was classified as

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TABLE 1—Distribution of Male Primary Cardiac Arrest Case-Control Pairs According to the Education of Wife: King County, Washington, December 1979–January 1981

		Education of Wife of Control		
		≤12 years	>12 years	Total
Education of Wife of Case	≤12 years	34	42	76
	>12 years	27	30	57
	Total	61	62	133

less than or equal to 12 years (high school graduate or less) or greater than 12 years (beyond high school), yielding approximately equal numbers in each group. Other risk factors which were measured included physical activity (estimated by the Minnesota Leisure-Time Activity Questionnaire),[14] alcohol consumption, current smoking status, and current treatment for hypertension. Odds ratios (OR) were estimated by the conditional maximum-likelihood method using discordant pairs[15] and by logistic regression for matched case-control data.[16] The association between wife's education and PCA in the husband was examined in the absence and presence of other risk factors and by level of the husband's own education, to detect a status incongruity effect.

Results.

The distribution of case-control pairs by wives' education is shown in Table 1. A wife who went beyond high school was less common among cases of PCA (OR = 0.6, 95 per cent CI = 0.4, 1.1). An OR of similar direction, magnitude, and precision was observed for the education of the men themselves (Table 2), with the more educated men at lower risk (OR = 0.7, 95 per cent CI = 0.4, 1.2). When variables representing wife's education and husband's education were entered into a conditional logistic model, the estimated OR for wife's education and PCA changed to 0.7 (95 per cent CI = 0.4, 1.2). The inclusion of a product term combining wife's education and husband's education revealed an absence of interaction; at both levels of husband's education, the estimated OR for wife's education was 0.7. When three categories of wife's education were considered (≤ 12 years, 13-15 years, ≥ 16 years) the OR's with ≤ 12 years as the referent category were 0.7 and 0.6, respectively.

The association between wife's education and primary cardiac arrest in the husband was also examined simultaneously with several risk factors (Table 3). The husband's education and activity level confounded the association slightly. Adjusted for all the risk factors, the OR for wife's education was 0.8 (0.5, 1.3).

Discussion

These results offer some evidence against the hypothesis that wife's education is directly associated with coronary heart disease in the husband.[1] As indicated by the confidence intervals, chance cannot exclude a weak association; also, there could have been unmeasured confounders in our data. If our findings are correct, differences between this study and previous studies may account for the discrepancy. For example, the effect of wife's education may vary for varying forms of coronary heart disease. Men in this study may differ from men studied previously, e.g., if few had Type A behavior.[3,17] However, the proportion of Seattle men with white collar occupations (52 per cent) would not indicate a likely deficit of Type A behavior.[18–20] The Seattle men were born somewhat later (mean year of birth = 1922) than the men of Framingham,[1] Rancho Bernardo,[2] or the Western Collaborative Group Study[3]; possibly secular changes in the social meaning and implications of wife's education have occurred.

Status incongruity and coronary heart disease has been a controversial subject, particularly since status incongruity is not easily defined and analyzed.[21–24] We view status incongruity effects as distinct from effects of the constituent status variables,[25] i.e., the risk for men with more educated wives would represent more than just the expected influence of two separate variables: wife's education and husband's education. These analyses suggest that a less educated husband and a more educated wife do not form a synergistic combination to increase risk of primary cardiac arrest.

TABLE 2—Distribution of Male Primary Cardiac Arrest Case-Control Pairs According to the Education of the Case/Control: King County, Washington, December 1979–January 1981

		Education of Control		
		≤12 years	>12 years	Total
Education of Case	≤12 years	26	38	64
	>12 years	26	43	69
	Total	52	81	133

TABLE 3—Adjusted Odds Ratios Relating Education of Wife to Primary Cardiac Arrest in the Husband: King County, Washington, December 1979–January 1981

Covariate*	Odds Ratio†	95% Confidence Interval
None	0.6	(0.4, 1.1)
Husband's education	0.7	(0.4, 1.2)
Husband's education, smoking	0.7	(0.4, 1.3)
Husband's education, hypertension	0.6	(0.4, 1.1)
Husband's education, physical activity	0.8	(0.5, 1.4)
Husband's education, alcohol consumption	0.7	(0.4, 1.2)
Husband's education, smoking, hypertension, physical activity, alcohol consumption	0.8	(0.5, 1.3)

*The covariates are defined as follows: husband's education (≤12/>12); current smoking status (yes/no); history of treatment for hypertension (yes/no); regular physical activity (above the median/below); monthly alcohol consumption (1–29 drinks/<1 drink; ≥30 drinks/<1 drink).

†Odds ratios were defined to contrast (wife's education >12 years)/(wife's education ≤12 years) and were derived from conditional logistic regression models.

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Effects of the North Carolina Prematurity Prevention Program among Public Patients Delivering at New Hanover Memorial Hospital

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Abstract: Twelve per cent of the 847 women who delivered in one hospital prior to implementation of the North Carolina Prematurity Prevention Program had low-birthweight births compared with 9.5 per cent of the 748 women who delivered during the program. Controlling for known risk factors, both low- and very-low birthweight births among Whites (Odds Ratio 2.0 and 3.7 respectively) and very-low-birthweight births among Blacks (OR 2.9) were reduced. (*Am J Public Health* 1988; 78:1493–1495.)

Introduction

In 1984, there were 8.2 neonatal deaths per 1,000 live births in North Carolina compared with 7.3 neonatal deaths per 1,000 live births in the United States in 1983.[1] Because premature delivery and low birthweight are major contributing factors to neonatal mortality,[2,3] the North Carolina Department of Human Resources initiated a statewide program to reduce the risk of prematurity; in January 1985 the obstetric clinic at New Hanover Memorial Hospital implemented this program.

Adapted from programs developed by Papiernik, Creasy, and Hobel,[4–6] the North Carolina Prematurity Prevention Program includes three components: patient identification, patient education, and staff education. Patient identification includes the use of a standardized risk assessment to identify women at risk of preterm labor. Patient education includes intensive education for all patients both on the signs and symptoms of preterm labor and on the need to seek early treatment if preterm labor occurs. Staff education includes in-service training for all staff on the importance of responding promptly to patient complaints of signs and symptoms of preterm labor. Staff are advised to freely admit and observe women suspected to be in preterm labor and, when appropriate, to treat these women using tocolytic agents.

New Hanover Memorial Hospital is a 520-bed community hospital with a university medical center affiliation. It is a regional referral center serving seven counties in the southeastern coastal area of North Carolina. Within the hospital, an obstetric outpatient clinic is staffed by University of North Carolina Medical School faculty, obstetric residents, staff nurses, and County Health Department personnel. Women seeking prenatal care at this facility are mainly young, poorly educated, economically disadvantaged, live in rural as well as urban areas, and about half are Black.

The purpose of this study is to evaluate the effectiveness of the Prematurity Prevention Program in reducing low-birthweight births in the population of women seeking prenatal care at this clinic.

Requests for a single reprint should be sent to Deborah Covington, MS, Research Associate, Area Health Education Center, 2131 South 17th Street, Wilmington, NC 28402. Co-authors Carl, Churchill, and Daley are also affiliated with the Area Health Education Center; Dr. Daley is also with the New Hanover Memorial Hospital in Wilmington, and the Department of Obstetrics and Gynecology, School of Medicine, University of North Carolina at Chapel Hill. Ms. Cushing is with the New Hanover County Health Department, Wilmington, NC. This paper, submitted to the Journal December 17, 1987, was revised and accepted for publication April 27, 1988.

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TABLE 1—Exclusion Criteria for the Two Study Groups

	Program Group	Pre-Program Group
Total registered for prenatal care	854	[a]
Aborted, ectopic, or molar pregnancy	31	—
Moved out of region	18	—
Transferred to private physician	11	—
Total delivered[b]	794	910
No prenatal care	32[c]	39
Twin gestation	6	11
Stillbirth	6	11
Livebirth <454 grams	2	2
Total included in analysis	748	847

a) Data on clinic registrants were not available prior to the Prematurity Prevention Program.

b) Labor and Delivery room statistics show that there were 98 fewer clinic staff deliveries in 1986 than in 1984. This could account for the difference between the two groups in the number of deliveries.

c) These women registered for the program, but did not return for prenatal care.

Methods

The study population included 854 women registering for prenatal care at the New Hanover Memorial Hospital obstetric clinic from July 1, 1985 to June 30, 1986.* One hundred and six women were excluded for reasons outlined in Table 1. Data for the 748 women included in the study were obtained by medical record and by birth certificate review.

To obtain comparison data, we used hospital records and birth certificate information for 910 New Hanover Memorial Hospital obstetric clinic patients who delivered prior to initiation of the program, from July 1, 1983 to June 30, 1984.** Using delivery records, we selected only women who received all their prenatal care and were delivered by New Hanover Memorial Hospital clinic staff, but excluded women who aborted, had ectopic pregnancies, or were referred for high-risk deliveries. A total of 847 women were included in the pre-program study group.

We compared the two groups with regard to the percentage of women who had low-birthweight (<2500 grams) and very-low birthweight births (<1500 grams). We used a logistic regression model to determine the independent effect of the program on decreasing low- and very-low-birthweight births while controlling for known risk factors.

*Although the program began in January 1985, the first six months was a trial period. Therefore, we began collecting data six months after program start-up.

**Because staff training took place prior to program start-up, we allowed a six-month interval to control for staff awareness and training effects.

Results

Table 2 compares the two study groups with regard to low birthweight and selected risk factors for low birthweight. With the exception of births less than 2500 grams among Blacks, women in the program were less likely to have a low- or a very-low-birthweight birth than women in the pre-program group. In both racial groups, there were differences in available risk factors between the pre-program and program groups.

Controlling for the potential confounding effects of the risk factors listed in Table 2, a logistic regression model showed that women in the pre-program group were 1.3 times more likely to have a low-birthweight birth than women in the program and 3.2 times more likely to have a very-low-birthweight birth (Table 3). Stratifying the logistic regression analysis by race, we found that White women in the pre-program group were 2.0 times more likely than women in the program group to have a low-birthweight birth and 3.7 times more likely to have a very-low-birthweight birth. Black women in the pre-program group were 2.9 times more likely to have a very-low-birthweight birth than those in the program group. The only other factor contributing to low and very-low birthweight in the logistic regression model was age less than 18 years (OR 1.6, 95% CI=1.2, 2.1 and OR 2.1, 95% CI=1.1, 3.2, respectively).

Discussion

In the initial phase of the Prematurity Prevention Program, results suggest that the program was effective in reducing low- and very-low-birthweight births. We speculate that the reason there was little difference in the percentages of low-birthweight births from pre-program to program group for Black women was that very-low-birthweight births were being delayed by program intervention, but not necessarily being delayed to term. Very-low-birthweight babies tend to have the highest rates of mortality, morbidity, and medical care costs. Delaying very-low-birthweight births even to low

TABLE 2—Low Birthweight and Selected Risk Factors for Low Birthweight by Group and by Race

	Program Group			Pre-Program Group		
	Total[a]	White	Black	Total[b]	White	Black
Number of Births	748	365	374	847	369	468
Low Birthweight	%	%	%	%	%	%
<2500 grams	9.5	6.6	12.5	12.3	12.2	12.2
<1500 grams	0.8	0.6	1.1	2.7	2.2	3.0
Risk Factors						
Black	50.0	—	—	55.3	—	—
Unmarried	57.6	39.2	75.6	50.6	28.0	68.4
Under age 18	12.2	11.5	12.6	13.4	12.2	14.5
Fewer than 12 years of education	44.2	48.2	38.9	49.3	57.9	42.9
Initiated prenatal care after first trimester	77.8	73.2	82.3	80.7	77.7	83.1

a) Includes nine cases categorized racially as "other".

b) Includes eight cases categorized racially as "other" and two as "unknown".

TABLE 3—Relative Odds of Having a Low-Birthweight Birth and Very-Low-Birthweight Birth for Women in the Pre-Program Group by Race Controlling for Certain Risk Factors[a]

		Low Birthweight (<2500 grams)		Very Low Birthweight (<1500 grams)	
	Number of Cases[b]	Odds Ratio	95% CI	Odds Ratio	95% CI
Total[c]	1578	1.3	(1.0, 1.7)	3.2	(2.3, 4.1)
White	728	2.0	(1.5, 2.6)	3.7	(2.2, 5.2)
Black	841	1.0	(0.6, 1.4)	2.9	(1.8, 3.9)

a) Risk factors controlled in the logistic regression model include less than 18 years of age, fewer than 12 years of schooling, initiation of prenatal care after the first trimester of pregnancy, single marital status, and for the total, Black race. Complete data available from author.
b) Cases with missing data were excluded from the analysis.
c) Includes cases categorized racially as "other" or unknown.

birthweight status both decreases mortality and morbidity and reduces the cost of infant care and the drain on critical care nursery resources.

A limitation of our study design is the use of historical controls to evaluate the program. We could not use a prospective randomized design because the program mandated that all women registering for prenatal care in the institution be included in the program; for this reason, we could not use women from the area in and out of the program during the same time period.

Another limitation of our study may be in the selection process. The Prematurity Prevention Program provided us with sufficient identifying information on each clinic registrant with which to select the program group. We had no such information prior to the program and therefore had to use delivery records to select the pre-program group. From these delivery records, we selected only women who had prenatal care at New Hanover Memorial Hospital obstetric clinic and who were delivered by clinic staff. This selection process should yield comparable groups; however, a caveat concerning selection bias seems warranted.

Finally, even though we controlled for certain known risk factors in the analysis, the two groups may differ on other factors for which we could not control.

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Johnson Foundation Funds New Program to Enhance Care for Elderly

The Robert Wood Johnson Foundation has launched a new program intended to develop model hospital systems of inpatient geriatric care in the United States. The initial grant of $392,000 was recently awarded to the Veterans Administration Medical Center in Denver — the first of four pilot projects to be funded under a $1.6 million national initiative. The Denver-based pilot will utilize the VA Medical Center's expertise in geriatrics, and also involve the participation of several major hospital and health provider organizations in Colorado.

Many elderly lack access to physicians trained in geriatrics. This program is designed to develop models to help community physicians and hospitals provide better care for the growing numbers of elderly patients in need of special attention. Interdisciplinary teams of medical personnel will be formed to address the varied and interrelated needs of the high-risk elderly patients.

A major focus of the Denver project will be to:

- Identify, upon admission, those elderly patients who are most at risk of functional decline during their hospitalization;
- Develop clinical management strategies to prevent problems that may arise during the hospital stay of these high-risk patients;
- Educate and enhance awareness of the special needs of the elderly patient in each participation institution and in the local community.

For further information, contact the RWJ Foundation, College Road, P.O. Box 2316, Princeton, NJ 08543-2316.

New from nchs

Dental Care and Dental Health: NHIS

MARY GRACE KOVAR, DRPH, SUSAN JACK, MS, AND BARBARA BLOOM, MPA

In 1986, 57 per cent of individuals in the United States age two years or over had made at least one visit to a dentist within a year compared with 55 per cent in 1983.

Receiving dental care is strongly associated with the ability to pay for it. People in families with more income were more likely to visit a dentist than people in families with less income; 74 per cent of the individuals in families with annual incomes of $35,000 or more had at least one dental visit in 1986, in contrast with 41 per cent in families with incomes of less than $10,000.

Having private dental insurance was associated with the use of dental care regardless of income. Individuals and families with private dental insurance were more likely to visit a dentist (70 per cent) than those without insurance (50 per cent). Those with private dental coverage had higher visit rates (2.6 versus 1.7 visits per person per year in 1986) and a higher proportion reported that their last visit was for a checkup. Such differences between those with and without private dental insurance persisted within each income category.

This information is from a recent report based on data from the 1986 National Health Interview Survey (NHIS).[1] It is the first time questions on dental care and dental health had been asked on that survey since 1983. Some questions were included from earlier surveys to examine change over time; new questions addressed such matters as dental insurance and children's use of prophylaxis.

The NHIS is a continuing household interview survey of the civilian noninstitutionalized population in the United States. US Bureau of the Census interviewers visit a probability sample of households each week. In 1986, 23,838 households containing 62,052 persons were interviewed. The household response rate was 96.5 per cent: most of the noninterviews were because no one was at home. A responsible member of the household answered the questions for children and for other adults in the family who were not at home.

Although access to dental care appears to have improved between 1983 and 1986, inability to pay, lack of awareness about the importance of dental care, or reluctance to seek dental care leave many people without adequate care.

Elementary school-age children need good dental care and 71 per cent of the children ages 5–11 years had seen a dentist at least once during the year prior to the interview. However, only 39 per cent had made two or more visits to a dentist. Over one-fourth of the children in this age span had not seen a dentist within a year, including 11 per cent who had never seen a dentist (Table 1). Forty per cent or more of the Black, Mexican American, or poor children had inadequate care (defined as no visit within a year). If inadequate is defined as no visit within six months, the percentages are much higher; only 44 per cent of the children 5–11 years of age were reported to have had a visit that recently. Even in families with incomes of $35,000 or more, 13 per cent of the children this age had not seen a dentist within a year and 39 per cent had not seen a dentist within six months.

TABLE 1—Children Ages 5–11 Years with Inadequate Dental Care

	No Visit	
	Within a year	Ever
	Per cent of Children	
All Children	28.3	11.4
Family Income less than $10,000	43.4	17.4
Mexican American	49.4	24.3
Black	41.3	16.9

SOURCE: 1986 National Health Interview Survey

According to their parents, 93 per cent of the children ages 2–16 used fluoride toothpaste, 8 per cent used fluoride supplements, and 13 per cent used a fluoride mouthrinse at home. About 7 per cent (12 per cent of the 9–11 year-olds) had received dental sealants.

About 11 per cent of the children were reported to be in fluoride mouthrinse programs at school, although the percentage of elementary school-age children in such programs is higher (Table 2). These programs are designed to serve low-income children and the data show that low-income

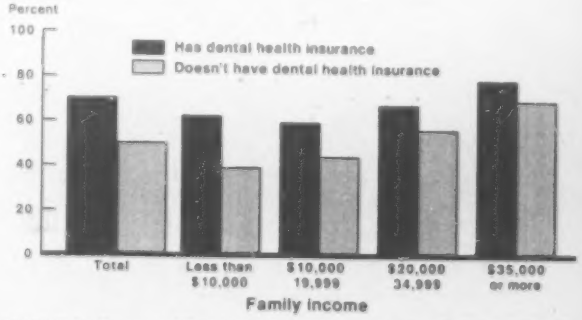

FIGURE 1—Per cent of Population Age 2 Years and Over with a Dental Visit within the Past Year by Family Income and Dental Insurance: United States, 1986

Copies of "Use of Dental Services and Dental Health" by S. Jack and B. Bloom, *Vital and Health Statistics* Series 10, No. 165 are available from the US Government Printing Office, Washington DC 20402. Order stock no. 017-022-0149-6.

For more information about this and other reports and data from the National Center for Health Statistics, write to the Scientific and Technical Information Branch, NCHS, 3700 East-West Highway, Hyattsville MD 20782, or call (301) 436-8500.

TABLE 2—Children's Participation in School Mouthrinse Programs

Age of Child (years)	Family Annual Income				
	All Incomes	Under $10,000	$10,000 –19,999	$20,000 –34,999	$35,000 or more
	Per Cent of Children				
2–16	10.5	13.5	12.8	10.7	7.5
5–8	15.5	18.8	18.6	15.8	11.2
9–11	19.8	22.4	24.4	20.2	15.3

SOURCE: 1986 National Health Interview Survey

children were more likely than upper-income children to be participating in such programs. However, only a minority of the children in low-income families were reported to be participating in fluoride mouthrinse programs at school. The report does not show whether those children had seen a dentist.

In contrast with children and younger adults, relatively few older people receive dental care. Only 42 per cent of the people age 65 and over had seen a dentist within a year; the critical distinction was whether the older person had any natural teeth. Among older people with teeth, 61 per cent had seen a dentist within a year; only 9 per cent of the older people with no remaining natural teeth had done so. The majority (71 per cent) of the edentulous older people had not seen a dentist for at least five years.

Dental care is an important, but sometimes neglected, part of total health care. People seldom die for lack of dental care, but the quality of their lives can be compromised by the lack of appropriate dental care. Children who fail to have their problems taken care of grow into adults with decayed, missing, or malocclude teeth. Older people who have not received appropriate care may suffer from periodontal disease, missing teeth, or poorly fitting dental appliances.

While the data in the report show that most people do receive dental care, they also document that significant groups of people do not; this can be useful for designing programs to reach those people.

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Kellogg Foundation Invites Applications for 1989 Fellowship Awards

The W. K. Kellogg Foundation recently announced the availability of its Kellogg National Fellowship awards for 1989. Initiated in 1980, the program is designed to prepare leaders who can function effectively and knowledgeably in dealing with complex problems where narrow expertise is not sufficient. It seeks to involve professional men and women in the earlier years of their careers who are interested in developing interdisciplinary and cross-cultural perspectives on contemporary human and social problems.

The Fellowship awards will be made on behalf of up to 50 individuals of exceptional merit and competence who have exhibited leadership potential in their community, organization or profession. To be eligible for the three-year program, an applicant must be a US citizen, accept responsibility to participate in *all* activities related to the program, and receive 25 per cent released time from his/her employer to carry out a non-degree, self-directed learning plan to expand personal horizons beyond the confines of a chosen profession. Individuals employed in both the non-profit and for-profit sectors may apply.

In addition to completion of the learning plan, Fellows must attend seven seminars sponsored by the Foundation, one of which is a two-week seminar in Latin America, with all others being one-week seminars in various cities in the US. Attendance is required at *all* seminars.

The award includes $35,000 for the development and completion of the learning plan and $5,000 to be used only for travel expenses incurred while participating in required seminars. In addition, 12-1/2 per cent of the Fellow's salary, up to $24,000 over the three years, will be reimbursed to the sponsoring/employing institution. Fellows not employed by a non-profit organization are not eligible for the salary reimbursement; however, in certain cases a monthly stipend of $300 may be available.

The deadline for applications is December 19, 1988. To obtain an application package or additional information, contact: Applications, Kellogg National Fellowship Program, W. K. Kellogg Foundation, 400 North Avenue, Battle Creek, MI 49017-3398. Tel: (616) 968-1611.

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PEDICULICIDAL/OVICIDAL ACTIVITIES: *In vitro* data indicate that permethrin has pediculicidal and ovicidal activity against *Pediculus humanus* var. *capitis*. The high cure rate (97-99%) of Nix in patients with head lice demonstrated at 14 days following a single application is attributable to a combination of its pediculicidal and ovicidal activities and its residual persistence on the hair which may also prevent reinfestation.

INDICATIONS AND USAGE: Nix is indicated for the single-application treatment of infestation with *Pediculus humanus* var. *capitis* (the head louse) and its nits (eggs). Retreatment for recurrences is required in less than 1% of patients since the ovicidal activity may be supplemented by residual persistence in the hair. If live lice are observed after at least seven days following the initial application, a second application can be given.

CONTRAINDICATIONS: Nix is contraindicated in patients with known hypersensitivity to any of its components, to any synthetic pyrethroid or pyrethrin, or to chrysanthemums.

WARNING: If hypersensitivity to Nix occurs, discontinue use.

PRECAUTIONS:

General: Head lice infestation is often accompanied by pruritus, erythema, and edema. Treatment with Nix may temporarily exacerbate these conditions.

Information for Patients: Patients with head lice should be advised that itching, redness, or swelling of the scalp may occur after application of Nix. If irritation persists, they should consult their physician. Nix is not irritating to the eyes; however, patients should be advised to avoid contact with eyes during application and to flush with water immediately if Nix gets in the eyes. In order to prevent accidental ingestion by children, the remaining contents of Nix should be discarded after use.

Combing of nits following treatment with Nix is not necessary for effective treatment. However, patients may do so for cosmetic or other reasons. The nits are easily combed from the hair treated with Nix after drying.

Carcinogenesis, Mutagenesis, Impairment of Fertility: Six carcinogenicity bioassays were evaluated with permethrin, three each in rats and mice. No tumorigenicity was seen in the rat studies. However, species-specific increases in pulmonary adenomas, a common benign tumor of mice of high spontaneous background incidence, were seen in the three mouse studies. In one of these studies there was an increased incidence of pulmonary alveolar-cell carcinomas and benign liver adenomas only in female mice when permethrin was given in their food at a concentration of 5000 ppm. Mutagenicity assays, which give useful correlative data for interpreting results from carcinogenicity bioassays in rodents, were negative. Permethrin showed no evidence of mutagenic potential in a battery of *in vitro* and *in vivo* genetic toxicity studies. Permethrin did not have any adverse effect on reproductive function at a dose of 180 mg/kg/day orally in a three-generation rat study.

Pregnancy: *Teratogenic Effects:* Pregnancy Category B: Reproduction studies have been performed in mice, rats, and rabbits (200-400 mg/kg/day orally) and have revealed no evidence of impaired fertility or harm to the fetus due to permethrin. There are, however, no adequate and well-controlled studies in pregnant women. Because animal reproduction studies are not always predictive of human response, this drug should be used during pregnancy only if clearly needed.

Nursing Mothers: It is not known whether this drug is excreted in human milk. Because many drugs are excreted in human milk and because of the evidence for tumorigenic potential of permethrin in animal studies, consideration should be given to discontinuing nursing temporarily or withholding the drug while the mother is nursing.

Pediatric Use: Nix is safe and effective in children two years of age and older. Safety and effectiveness in children less than two years of age have not been established.

ADVERSE REACTIONS: The most frequent adverse reaction to Nix is pruritus. This is usually a consequence of head lice infestation itself, but may be temporarily aggravated following treatment with Nix. 5.9% of patients in clinical studies experienced mild temporary itching; 3.4% experienced mild transient burning/stinging, tingling, numbness, or scalp discomfort; and 2.1% experienced mild transient erythema, edema, or rash of the scalp.

DOSAGE AND ADMINISTRATION:

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Patterns of Sexual Behavior and Condom Use in a Cohort of Homosexual Men

MARTIN T. SCHECHTER, MD, PHD, KEVIN J. P. CRAIB, MMATH, BRIAN WILLOUGHBY, MD, CCFP, BRUCE DOUGLAS, MD, FRCPC, W. ALASTAIR MCLEOD, MD, FRCPC, MICHAEL MAYNARD, MD, PETER CONSTANCE, MD, AND MICHAEL O'SHAUGHNESSY, PHD

Abstract: To measure the magnitude of risk reduction within a cohort of homosexual men, questionnaire responses in April 1984–March 1985 were compared to those in October 1986–September 1987. A total of 361 subjects were eligible (231 seronegative, 130 seropositive). The annual number of sex partners declined significantly from a median of 8.0 to 5.0 with no difference between the serologic groups. The number of subjects reporting no receptive anal intercourse increased as did condom use during anal receptive intercourse. More seronegatives than seropositives subjects reported *no* condom use during receptive anal intercourse with *regular* partners (45.7 per cent versus 23.4 per cent), and with *casual* partners (14.9 per cent versus 1.5% percent). Among subjects with the most casual sexual contact at the second visit, 33.3 per cent of seronegatives and 29.2 per cent of seropositives did not report usual condom use during receptive anal intercourse with casual partners. Although we have documented marked risk reductions, safe sex practices are still not universal, and a few individuals continue to put themselves at extremely high risk. (*Am J Public Health* 1988; 78:1535–1538.)

Introduction

In the absence of an effective vaccine against *human immunodeficiency virus* (HIV), primary prevention efforts must center on modifying behaviors associated with its transmission. Thus, community groups and public health agencies have concentrated their efforts to varying degrees on promoting both reductions in the number of sex partners and in the frequency of anal intercourse, as well as increases in the use of condoms and in the practice of other safe sexual activities. The degree to which such changes have occurred has only recently been assessed and the effect of knowledge of HIV serologic status on risk reduction remains unclear. We report here on changes in sexual practices and condom use in our ongoing cohort study of homosexual men.

Methods

The methods and aims of the Vancouver Lymphadenopathy-AIDS Study (VLAS) have been described in detail previously.[1,2] Briefly, the VLAS is an ongoing prospective study of over 700 homosexual men who were recruited from six general practices in central Vancouver, British Columbia during the period November 1982 to February 1984. Subjects have been returning for follow-up visits approximately every six months since enrollment. The visit frequency has been reduced to once per year in the last cycle. During each visit, subjects complete a detailed self-administered questionnaire. In addition, each study subject undergoes a complete physical examination and blood samples are drawn for immunologic and HIV antibody testing. Antibody testing is carried out at the National Reference Laboratory at the Laboratory Center for Disease Control in Ottawa utilizing ELISA (enzyme linked immunosorbent assay) with equivocal results further assessed by Western Blot. Some participants may have additional blood samples drawn for HIV antibody testing between study visits at the discretion of their physicians. All participants are counseled by their physicians to adopt safer sex guidelines including reduction in the number of sex partners and avoidance of unprotected intercourse.

Subjects were eligible for this analysis if they completed their third visit during the period April 1984–March 1985 (EV) and their seventh visit during October 1986–September 1987 (LV). The third visit was chosen for the baseline because serologic results were not available until this time and the seventh visit was the most recently completed at the time of analysis.

As of October 1987, 58 cases of AIDS (acquired immune deficiency syndrome) had been diagnosed in our study according to Centers for Disease Control (CDC) criteria. Seropositive subjects for the present analysis were defined as those eligible subjects who were seropositive at EV and who had remained AIDS-free as of October 1987. It should be noted, however, that 14 per cent of the seropositive group at EV had an AIDS-related condition as defined by a helper cell (CD4) count below 400 and/or two or more of the following symptoms: fever, weight loss, fatigue, diarrhea, arthralgia, thrush, cough unrelated to smoking, or dyspnea. Seronegative subjects were defined as those who were seronegative at EV. It was recognized that the seronegative group would contain individuals who seroconverted between visits and were therefore seropositive at the time of LV.

Patterns of sexual behavior were measured using questionnaire items pertaining to the number of different sexual partners in the previous 12 months, the frequency of various sexual practices during this period, and condom use during these practices. At LV, a distinction was made between regular (at least one sexual encounter per month) and casual partners.

Statistical analyses of the data were carried out using non-parametric methods because inspection revealed the distribution of number of sexual partners to be non-normal. All reported p-values are two-sided. For all paired comparisons, subjects with incomplete or missing information regarding the variable of interest were excluded from the analysis.

Results

A total of 361 subjects completed their third and seventh

Address reprint requests to Martin T. Schechter, MD, PhD, Vancouver Lymphadenopathy-AIDS Study Group, St. Paul's Hospital, 1081 Burrard Street, Vancouver, British Columbia, Canada V6Z 1Y6. He and all the co-authors (except Dr. O'Shaughnessy) are with the AIDS Study Group and also affiliated with the University of British Columbia; Dr. O'Shaughnessy is with the Viral Laboratory, Federal Center for AIDS, Department of National Health and Welfare, Ottawa, Canada. This paper, submitted to the Journal February 22, 1988, was revised and accepted for publication May 18, 1988.

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visits within the defined time periods. Those not eligible included 97 whose third and/or seventh visits fell outside the defined periods, 45 who had not yet completed their seventh visit, 58 who had developed AIDS by October 1987, 79 who had moved away, 89 who had withdrawn from the study, and one who died of unrelated causes. Of the 361 eligible subjects, 231 (64.0 per cent) were seronegative at EV, and 130 (36 per cent) were seropositive; 27 (11.7 per cent) of the 231 seronegative subjects seroconverted during the period between visits. The median interval duration between the EV and LV for all 361 subjects was 28 months (range = 20–42), 29 months for the seronegative group, and 27 months for the seropositive group. The median ages at EV were identical (34 years) for the two groups.

As seen in Table 1, there was a decline in the reported number of annual sexual partners between the two visits from a median of 8.0 to 5.0. Of the 361 subjects, 246 (68.1 per cent) reported a reduction in the number of sexual partners, 19 (5.2 per cent) reported no change, and 96 (26.6 per cent) reported an increase. Declines were noted in both the seronegative and seropositive groups. It is noteworthy that the decline in numbers of partners was similar whether or not an AIDS-related condition as defined above was present at EV.

To control for the higher number of sexual partners at baseline in the seropositives, we restricted the analysis to subjects who reported more than the median number (8) of sexual partners at EV (Table 2). Essentially equal declines were noted in both groups.

As seen in Table 3, there was a decline in the frequency of receptive anal intercourse between EV and LV. At EV, 28.1 per cent of seronegatives and 16.9 per cent of seropositives reported none of this practice in the prior year. At LV, these proportions had risen to 45.3 per cent and 35.6 per cent, respectively with regular sexual partners and 74.5 per cent and 36.9 per cent, respectively with casual partners. However, 6.6 per cent of seronegatives and 31.1 per cent of seropositives reported receptive anal intercourse in at least 30 per cent of casual sexual encounters. Seronegatives were much more likely to not practice receptive anal intercourse with casual than with regular partners at LV but no such differences were seen with seropositives.

Table 4 demonstrates a marked increase between the visits in the use of condoms during receptive anal intercourse. At EV, only 3.6 per cent of seronegatives and 3.7 per cent of seropositives reported condom use in more than 30 per cent of receptive anal acts. At LV, these proportions had risen to 34.4 per cent and 63.9 per cent, respectively with regular partners and 72.3 per cent and 81.5 per cent, respectively with casual partners. It is noteworthy that a higher proportion of seronegative subjects than seropositive subjects reported never using condoms during receptive anal intercourse with casual partners (14.9 per cent vs 1.5 per cent; $p = 0.007$).

As seen in Table 5, among those in the upper 50 per cent of casual contact, 33.3 per cent of seronegatives and 28.2 per cent of seropositives reported they used condoms in less than 60 per cent of receptive anal exposures to casual partners.

TABLE 1—Reported Numbers of Sexual Partners in Past Year at EV and LV by Serologic Group

Group	EV	LV	LV-EV*	P-Value**
Seronegative (n = 231)				
Median	6.0	5.0	−2.0	<.0001
Range	[0,140]	[0,208]	[−99, +68]	
Mean	16.1	12.2	−3.9	
SD	23.2	20.6	17.6	
Seropositive (n = 130)				
Median	12.0	7.0	−3.0	<.0001
Range	[2,200]	[0,100]	[−151, +88]	
Mean	21.0	14.0	−7.0	
SD	26.1	17.9	27.0	
All subjects (n = 361)				
Mean	8.0	5.0	−2.0	<.0001
Range	[0,200]	[0,208]	[−151, +88]	
Mean	17.9	12.8	−5.1	
SD	24.3	19.7	21.5	

*Paired difference LV-EV
**P-value based on Wilcoxon signed rank test.

TABLE 2—Reported Numbers of Sexual Partners in Past Year at EV and LV by Serologic Group (restricted to those reporting more than eight sexual partners at EV)

Group	EV	LV	LV-EV*	P-Value**
Seronegative (n = 94)				
Median	24.0	15.0	−9.0	<.0001
Range	[10,140]	[1,208]	[−99, +68]	
Mean	33.5	22.0	−11.5	
SD	28.4	28.0	24.0	
Seropositive (n = 78)				
Median	24.0	12.0	−10.0	<.0001
Range	[10,200]	[1,100]	[−151, +70]	
Mean	32.0	18.1	−13.9	
SD	28.8	18.5	30.7	
All subjects (n = 172)				
Median	24.0	12.0	−9.0	<.0001
Range	[10,200]	[1,208]	[−151, +70]	
Mean	32.8	20.3	−12.5	
SD	28.5	24.2	27.1	

*Paired difference of LV-EV
**P-value based on Wilcoxon signed rank test.

Discussion

There are several caveats which should be recognized at the outset concerning these data. Our original intent in recruiting participants through general practitioners was to avoid potential bias associated with recruitment through sexually transmitted disease clinics or secondary referral centers. Nevertheless, it must be recognized that gay men who choose to attend medical practices that provide care to

TABLE 3—Frequency of Receptive Anal Intercourse by Serologic Status at EV and LV

Sexual Behavior	Frequency*	Seronegative n (%)	Seropositive n (%)
Receptive anal intercourse at EV	Never	64 (28.1)	22 (16.9)
	Infrequent	104 (45.6)	51 (39.2)
	Often	34 (14.9)	28 (21.5)
	Usually	26 (11.4)	29 (22.3)
Receptive anal intercourse with regular partners at LV**	Never	58 (45.3)	26 (35.6)
	Infrequent	47 (36.7)	28 (38.4)
	Often	9 (7.0)	7 (9.6)
	Usually	14 (10.9)	12 (16.4)
Receptive anal intercourse with casual partners at LV†	Never	137 (74.5)	38 (36.9)
	Infrequent	35 (19.0)	33 (32.0)
	Often	6 (3.3)	17 (16.5)
	Usually	6 (3.3)	15 (14.6)

*Infrequent = 1–30% of sexual encounters; Often = 31–60% of sexual encounters; Usually = >60% of sexual encounters
**Restricted to those reporting at least one regular partner
†Restricted to those reporting at least one casual partner

TABLE 4—Frequency of Condom Use During Receptive Anal Intercourse at EV and LV by Serologic Group Among Those Reporting Receptive Anal Intercourse

Sexual Behavior	Frequency of Condom Use*	Seronegative** n (%)	Seropositive** n (%)
Receptive anal intercourse at EV	Never	130 (79.3)	78 (72.2)
	Infrequent	28 (17.1)	26 (24.1)
	Often	3 (1.8)	3 (2.8)
	Usually	3 (1.8)	1 (0.9)
Receptive anal intercourse with regular partners at LV**	Never	32 (45.7)	11 (23.4)
	Infrequent	14 (20.0)	6 (12.8)
	Often	2 (2.9)	6 (12.8)
	Usually	22 (31.4)	24 (51.1)
Receptive anal intercourse with casual partners at LV†	Never	7 (14.9)	1 (1.5)
	Infrequent	6 (12.8)	11 (16.9)
	Often	4 (8.5)	8 (12.3)
	Usually	30 (63.8)	45 (69.2)

*Infrequent = 1–30% of episodes of receptive anal intercourse; Often = 31–60% of episodes of receptive anal intercourse; Usually = >60% of episodes of receptive anal intercourse.

**Restricted to those reporting at least some receptive anal intercourse with a regular partner.

†Restricted to those reporting at least some receptive anal intercourse with a casual partner.

large numbers of gay men may not accurately reflect the homosexual community at large. Furthermore, the mere participation in a prospective study involving repeated questionnaires, physical examinations, serologic testing, and counseling by committed practitioners can be expected to have additional effects on behavior that may not have occurred outside the context of the study. This represents a form of surveillance bias in which the very act of observation influences the behavior being observed. In addition, questions naturally arise concerning the reliability and validity of information obtained by questionnaire regarding sexual activity. However, some reports suggest that both the interviewer-administered and self-administered questionnaire can provide reasonably reliable data concerning sexual behavior in gay men when the period of recall is relatively short.[3,4] The data analyzed in this study always pertained to sexual behavior in at most the year prior to questionnaire completion. There remains the question of the definition of the seronegative group. We defined the latter to include those individuals who were HIV antibody negative at the time of EV which resulted in the inclusion of 27 individuals who had seroconverted between EV and LV. However, as it was our intent to study the relation of serologic status and *subsequent* behavior change, we felt that this definition was the most appropriate. Any definition requiring persistent seronegativity throughout the duration of observation would impose constraints on behavioral variables during this period and artifactually influence the behavior change under study.

TABLE 5—Frequency of Condom Use During Receptive Anal Intercourse with Casual Partners at LV by Serologic Group Among Those Reporting Receptive Anal Intercourse and More Than Five Casual Partners

Sexual Behavior	Frequency of Condom Use*	Seronegative** n (%)	Seropositive** n (%)
Receptive anal intercourse with casual partners at LV	Never	3 (10.0)	0 (0.0)
	Infrequent	4 (13.3)	8 (16.7)
	Often	3 (10.0)	6 (12.5)
	Usually	20 (66.7)	34 (70.8)

*Infrequent = 1–30% of episodes of receptive anal intercourse; Often = 31–60% of episodes of receptive anal intercourse; Usually = >60% of episodes of receptive anal intercourse.

The present data document marked changes in sexual behavior in this cohort over a 28-month period, suggesting at least some success for ongoing educational initiatives. Our results are consistent with behavioral change which has been reported in several other studies of homosexual men[5–13] and which may now be having beneficial effects on HIV transmission. Winkelstein, *et al*, demonstrated progressive declines in seroconversion rates in a seronegative cohort which they attributed, at least in part, to behavior change.[5] We have recently found similar declines in seroconversion among initially seronegative men in our cohort, with five successive annual seroconversion rates from November 1982 through October 1987 of 3.9, 9.8, 7.3, 4.2, and 0.9 per cent respectively.*

Our finding that the decline in number of sexual partners was comparable in the seronegative and seropositive groups is consistent with the observation of Fox, *et al*,[12] that disclosure of serologic status did not markedly influence subsequent numbers of partners in a cohort of homosexual men. The fact observed here and elsewhere[12,13] that the marked overall behavior change exceeded any differences between serologic groups suggests that the effects of societal influences, such as community education and media coverage, may have outweighed any marginal additional effect of knowledge of serologic status. We did observe some differences, however, between the serologic groups. Although seronegatives practiced receptive anal intercourse with casual partners less often than did seropositives, it was disconcerting to note that when they did engage in this practice, seronegatives utilized condoms less often than did seropositives. Although it may be that seronegatives were more selective in the choice of their casual partners, the possibility remains as suggested by others[12,13] that some seronegatives may infer from their negative status some form of protection and this may act as a disincentive to behavior change.

Our data highlight an important fact concerning sexual behavior in this population often overlooked in studies of temporal change. It is that it is potentially misleading to measure overall change in certain variables such as numbers of partners or condom use without reference to the type of partner or the sexual activity in question. Sexual encounters can vary in their risk of HIV transmission from the ongoing sexual encounters of two mutually monogamous seronegative partners at one end of the spectrum, to receptive anal intercourse with casual or anonymous partners of unknown serology at the other. Seronegatives appeared to adjust the type of sexual practice to the partner in that they were much more likely to never practice receptive anal intercourse with casual partners than with regular partners. Furthermore, at the recent visit, seronegatives were more likely to use condoms during receptive anal intercourse with casual partners than with regular partners. Failure to distinguish between types of partners would have obscured these differences and would have attenuated the change observed with casual partners where safer practices are most critical. Nevertheless, it was disappointing to observe that even at the recent visit, as many as 15 per cent of seronegatives never use condoms during receptive anal intercourse with casual part-

* Schechter MT, Willoughby B, Craib KJP, *et al*. Unpublished data, 1988.

ners and only 64 per cent of this group report the usual use of condoms during this practice. While it is a small proportion of seronegatives who report continuing receptive anal intercourse with casual partners, these individuals continue to place themselves at extremely high risk of HIV infection and probably account for why new infections, although slowing, are still continuing to appear. Also, seropositives with symptoms or low helper cell counts, who may be the most infectious to others,[14] did not lower their numbers of partners to any greater degree than well seropositives. The barriers to behavior change[15,16] at work in these subgroups and in other areas where behavior change has been less marked[17,18] need to be studied so that individuals at continuing high risk can be effectively targeted.

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US Government Issues Report on Status of Organ Transplantation

The US Government issued its first annual report on the scientific and clinical status of organ transplantation in late summer. The essence of the report, based on the 1986 data, is that "significant advances are being made in the field of organ transplantation," and there has been a dramatic rise in the number of people receiving transplants as well as the number of people on waiting lists. According to the report:

- One-year patient survival is 96 per cent for recipients of kidneys from living donors, and 91 per cent for recipients of cadaveric kidneys;
- One-year graft survival rates are about 88 per cent for recipients of kidneys from living donors, and 71 per cent for recipients of cadaveric kidneys;
- The probability of survival for at least one year after heart transplantation is about 85 per cent when triple drug therapy is used to suppress the immune system. Triple drug therapy usually involves cyclosporine, azathioprine and corticosterioids.
- The number of people benefitting from organ transplants—kidney, heart, liver, heart-lung, or pancreas transplants—was about 11,500 in 1986. More than 10,500 remained on waiting lists at the end of the year.

Copies of the new *Report on the Scientific and Clinical Status of Organ Transplantation* may be obtained from: Division of Organ Transplantation, Rm 9-31, Health Resources and Services Administration, US Public Health Service, DHHS, 5600 Fishers Lane, Rockville, MD 20857. Tel: (301) 443-7577.

Racial/Ethnic Differences in the Risk of AIDS in the United States

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Abstract: We analyzed the variation in the risk of AIDS in US Blacks, Hispanics, and other racial/ethnic groups relative to that in Whites (non-Hispanic) by geographic area and mode of acquiring HIV infection, based on data reported between June 1, 1981 and January 18, 1988 to the Centers for Disease Control and 1980 US census data. Relative risks (RRs) in Blacks and Hispanics were highest in the northeast region, and higher in suburbs than in central cities of metropolitan areas. RRs in Blacks and Hispanics were greatest for AIDS directly or indirectly associated with intravenous-drug abuse by heterosexuals (range: 5.7–26.9) and were also high for AIDS associated with male bisexuality (range: 2.5–4.8), suggesting that these behaviors may be more prevalent in Blacks and Hispanics than in Whites. Prevention strategies should take into account these racial/ethnic differences. (*Am J Public Health* 1988; 78:1539–1545.)

Introduction

Although most patients reported with acquired immunodeficiency syndrome (AIDS) in the United States are non-Hispanic White homosexual males, the perception of AIDS as predominantly a disease of gay White men has changed. The risk of AIDS is higher in American Blacks and Hispanics and the racial/ethnic differences in risk are greatest in association with intravenous-drug abuse (IVDA) by heterosexuals.[1–9] We present here the results of a study to determine more precisely the magnitude of the association between AIDS and racial/ethnic group for different means of acquiring human immunodeficiency virus (HIV) infection and the variation in this association by geographic area and over time.

Methods

To measure the risk of AIDS in a given racial/ethnic group, we used the cumulative incidence (number of AIDS cases reported in that group per million population of the same group). AIDS cases were those meeting the case definition[10] in residents of the United States (excluding US territories) reported to the Centers for Disease Control (CDC) between June 1, 1981 and January 18, 1988. Reporting of AIDS cases by health care providers is mandatory in all states. CDC receives data on these cases from the health departments of state and selected local governments. The population of each racial/ethnic group was that reported by the US Bureau of the Census for 1980.[11]

We used relative risk (RR) to examine the magnitude of the association between AIDS and racial/ethnic group. The RR in a given group was the ratio of the cumulative incidence of AIDS in that group to the cumulative incidence in a reference group, in this instance being Whites who are not Hispanic. We stratified the analyses by sex and age group (men, women, and children <15 years of age).

The 99 per cent confidence intervals around RRs were approximated by the first-order Taylor series method.[12] For RRs equal to zero, the upper limits of the confidence intervals were estimated by the method of Fleiss.[13]

We classified AIDS patients in a hierarchy of exposure categories according to their presumed means of acquiring HIV infection. Patients with more than one possible means of acquisition were classified only in the exposure category listed first in the hierarchy, except for the combination of IVDA and male homosexuality/bisexuality.

We also stratified the analyses by region,[11] population of metropolitan areas, and residence inside or outside the central cities of selected metropolitan areas. For analysis by residence inside or outside the central cities of metropolitan areas, we selected the 16 most populous SMSAs (≥2 million inhabitants each) that had central cities.

To assess trends in RRs by exposure category from 1979 through 1987, we used weighted least-squares regression analysis (with the inverse variances of the quarter-yearly RRs as weights).[15] Only RRs significant at the $p<0.01$ level were analyzed.

Results

CDC received reports of 50,830 AIDS patients in the United States between June 1, 1981 and January 18, 1988. Of the 50,704 (99.8 per cent) with specified race/ethnicity, 60.9 per cent were classified as White, 25.5 per cent as Black, 12.9 per cent as Hispanic, and 0.7 per cent as Other. In contrast, of the US population (226,545,805), 79.6 per cent were classified as White, 11.5 per cent as Black, 6.4 per cent as Hispanic, and 2.5 per cent as Other. Overall, the risk (cumulative incidence) of AIDS in Blacks and Hispanics was almost three times as great as that in Whites. By sex and age group, the risks of AIDS in Black and Hispanic men were 2.8 and 2.7 times, respectively, that in White men; the risks in Black and Hispanic women were 13.2 and 8.1 times that in White women; and the risks in Black and Hispanic children were 11.6 and 6.6 times that in White children (Table 1). The risks in other racial/ethnic groups were less than the risk in Whites.

Analysis by Exposure Category

The proportion of AIDS cases associated with IVDA by heterosexuals (including their sex partners and children) was 39.7 per cent in Blacks and 38.7 per cent in Hispanics, compared with 6.3 per cent in Whites, 4.0 per cent in Asians and Pacific Islanders, and 19.6 per cent in American Indians, Eskimos, and Aleuts. These IVDA-exposure categories were associated with the highest RRs in Blacks and Hispanics, ranging from 5.7 to 26.9 (Table 2).

The risk of AIDS in exclusively homosexual men without a history of IVDA was 1.3 and 1.7 times as great in Black and Hispanic men, respectively, as in White men (Table 3). For bisexual men with AIDS, the RRs were higher (3.6 and 2.5 in Blacks and Hispanics, respectively). The proportion of bisexual men among homosexual and bisexual men with

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TABLE 1—Cumulative Incidence* and Relative Risk (RR)† of AIDS, by Racial/Ethnic Group

Sex and Age Group	Racial/Ethnic Group	AIDS Cases	Cumulative Incidence*	RR	(99% Confidence Interval)
Men,					
≥15 years	White	29,614	432.9	1.0	
	Black	10,569	1,233.7	2.8	(2.8, 2.9)
	Hispanic	5,783	1,179.7	2.7	(2.6, 2.8)
	Other	317	123.1	0.4	(0.3, 0.4)
Women,					
≥15 years	White	1,090	14.7	1.0	
	Black	1,946	193.4	13.2	(12.0, 14.5)
	Hispanic	596	118.5	8.1	(7.1, 9.2)
	Other	26	12.6	0.9	(0.5, 1.4)
Children,					
<15 years	White	183	4.9	1.0	
	Black	423	56.6	11.6	(9.3, 14.6)
	Hispanic	151	32.3	6.6	(5.0, 8.8)
	Other	6	3.9	0.8	(0.3, 2.3)
Total	White	30,887	171.4	1.0	
	Black	12,938	495.6	2.9	(2.8, 3.0)
	Hispanic	6,530	447.0	2.6	(2.5, 2.7)
	Asian or Pacific Islander	298	83.9	0.5	(0.4, 0.6)
	American Indian, Eskimo, or Aleut	51	35.6	0.2	(0.1, 0.3)

*AIDS cases reported from June 1, 1981 to January 18, 1988 per million population of same race/ethnicity.
†Reference group for relative risk: Non-Hispanic Whites

AIDS was greater in Blacks (30 per cent) and Hispanics (20 per cent) than in Whites (14 per cent). This pattern was reflected in the risk of AIDS in women whose sex partner was a bisexual man, which was 4.6 and 3.6 times as great in Black and Hispanic women, respectively, as in White women. The addition of a history of IVDA increased the racial/ethnic RRs associated with male homosexuality (Table 3).

The risk of AIDS associated with prior receipt of a blood transfusion was 2.5 times as great in Black and Hispanic children as in White children. In addition, the risks of AIDS in transfused women, in male sex partners of transfused women, and in children of transfused women were higher in Blacks than in Whites (Table 4). The only category in which Blacks had a lower risk of AIDS than Whites was coagulation disorders (e.g., hemophilia), in which the RR in Black men was 0.4 (99 per cent confidence interval: 0.3, 0.8).

The risks of AIDS in Blacks and Hispanics were substantially higher than the risk in Whites for exposure categories associated with undetermined means of acquiring HIV infection, including categories involving heterosexual and perinatal transmission from persons with undetermined means of acquiring HIV infection (Table 5).

Analysis by Geographic Area

In general, the RRs in Blacks and Hispanics were greatest in the northeast US and least in the west (Table 6). In the west, the risk of Hispanics was actually less than that in Whites for AIDS in exclusively homosexual men and for AIDS overall. In all regions, the risk in other groups (e.g., Asians) was less than the risk in Whites for AIDS in homosexual men and overall; for no exposure category was it significantly greater than the risk in Whites. The northeast was the only region in which most AIDS cases in Blacks and Hispanics (54 per cent and 57 per cent, respectively) were associated with IVDA by heterosexuals (including their sex partners and children).

Of AIDS patients with a reported city or county of residence (97 per cent), 95 per cent resided in 314 of the 318 SMSAs recognized by the Bureau of the Census; 5 per cent resided outside SMSAs. The 168 SMSAs in the smallest population category (50,000–250,000 population) contained 10 per cent of the US population and accounted for 3 per cent of AIDS cases, whereas the 16 SMSAs in the largest population category (>2,000,000 population) contained 27 per cent of the US population and accounted for 62 per cent

TABLE 2—Relative Risk* of AIDS Associated with History of Intravenous-Drug Abuse (IVDA), by Racial/Ethnic Group and Exposure Category

Exposure Category	White*	Black	Hispanic	Other
Adults with IVDA:				
Homosexual or bisexual men	1.0	3.0 (2.7, 3.3)	2.3 (2.0, 2.6)	0.2 (0.1, 0.4)
Heterosexual men	1.0	20.7 (19.1, 22.6)	20.6 (18.8, 22.6)	0.2 (0.1, 0.6)
Women	1.0	18.1 (15.7, 21.0)	10.3 (8.5, 12.5)	0.7 (0.3, 1.7)
Adults whose sex partner had IVDA:				
Heterosexual men	1.0	15.3 (9.8, 23.9)	5.7 (2.9, 11.2)	0.7 (0.1, 9.2)
Women	1.0	21.9 (16.4, 29.2)	22.7 (16.5, 31.3)	0.3 (0.0, 4.5)
Children whose mother:				
Had IVDA	1.0	26.9 (17.0, 42.7)	16.7 (10.0, 28.0)	1.3 (0.2, 8.5)
Had sex partner with IVDA	1.0	18.7 (8.6, 40.5)	17.8 (7.8, 40.8)	0.0 (0.0, 9.7)

*Reference group: non-Hispanic Whites (99% confidence interval around relative risk is in parentheses).

TABLE 3—Relative Risk* of AIDS Associated with History of Male Homosexual Activity, by Racial/Ethnic Group and Exposure Category

Exposure Category	White*	Black	Hispanic	Other
Exclusively homosexual men:				
without IVDA†	1.0	1.3 (1.2, 1.4)	1.7 (1.6, 1.8)	0.3 (0.3, 0.4)
with IVDA	1.0	2.5 (2.2, 2.8)	2.0 (1.7, 2.3)	0.2 (0.1, 0.4)
Bisexual men:				
without IVDA	1.0	3.6 (3.3, 3.9)	2.5 (2.2, 2.8)	0.6 (0.4, 0.8)
with IVDA	1.0	4.8 (4.0, 5.7)	3.3 (2.6, 4.3)	0.1 (0.0, 0.8)
Women whose sex partner was a bisexual man	1.0	4.6 (2.9, 4.7)	3.6 (1.9, 6.9)	0.9 (0.1, 5.8)
Children whose mother's sex partner was a bisexual man	1.0	3.8 (0.9, 15.1)	1.0 (0.1, 15.4)	0.0 (0.0, 17.1)

*Reference group: non-Hispanic Whites (99% confidence interval around relative risk is in parentheses).
†History of intravenous-drug abuse.

of AIDS cases. The cumulative incidence of AIDS was correlated with the population of SMSAs ($p=0.0001$, Spearman rank sum correlation coefficient $=0.49$). The cumulative incidence was higher in Blacks and Hispanics than in Whites in each population category of SMSAs and outside SMSAs (RR range: 1.3–3.7) (Figure 1).

In the 16 SMSAs with ≥2,000,000 inhabitants, the cumulative incidence of AIDS was higher in the central cities than in the suburbs. For cases in exclusively homosexual men, in the central cities, the risk in Blacks and Hispanics was about half that in Whites; in the suburbs the risk in Blacks and Hispanics was slightly higher than the risk in Whites (Table 7). For cases associated with bisexual men, heterosexual intravenous drug abusers (IVDAs), and undetermined means of acquiring HIV infection, the risks in Blacks and Hispanics were higher than the risk in Whites in both the central cities and the suburbs. The pattern was similar in each of the four regions of the United States, except that, in the west (Los Angeles and San Francisco-Oakland), the risk of AIDS in exclusively homosexual men was lower in Blacks and Hispanics than in Whites in the suburbs as well as in the central cities.

Analysis by Time of Diagnosis

In 1980, the annual incidence of AIDS in heterosexual IVDAs was already 23.0 and 28.7 times as great in Blacks and Hispanics, respectively, as in Whites. By 1987, in Hispanics the RR of AIDS in heterosexual IVDAs had dropped to 14.8 (regression coefficient $=-0.38$, SE $=0.15$) along with a decline of Hispanics' overall RR of AIDS from 3.2 in 1981 to 2.5. The declining RR reflected an incidence increasing at a slower rate than that in Whites, rather than a decreasing incidence. The RR in Blacks has not had a significant downward or upward trend for AIDS overall or for any exposure category. In other racial/ethnic groups, the RR rose from 0.1 in 1982 to 0.4 in 1987, both overall and for cases in homosexual/bisexual men (regression coefficient $=0.018$, SE $=0.004$).

Discussion

The wide range in the RRs of AIDS in Blacks and Hispanics for different exposure categories, and the variation in these RRs by geographic area of residence, support the view that the higher risks of AIDS in Blacks and Hispanics are due primarily to behavioral and perhaps environmental differences between the racial/ethnic groups, rather than genetic differences. A large part of the difference in cumulative incidence between racial/ethnic groups may be due to different levels of prevalence of the various means of acquiring HIV infection, but reliable data on risk factor prevalence by racial/ethnic group are not currently available.

Analyses of data from drug treatment programs and emergency room visits for drug-related conditions are biased by geographical location and/or exclusion of data on private patients or drug abusers not seeking treatment.[9,16,17] Thus, the disproportionate numbers of Black and Hispanic clients in these settings provide suggestive but not conclusive evidence of a higher prevalence of IVDA in Blacks and Hispanics.

Studies of heterosexual IVDAs also have found a higher prevalence and incidence of HIV infection in Blacks and Hispanics than in Whites,[18–23] even after controlling for frequency of needle sharing.[18,20–22] This might have resulted

TABLE 4—Relative Risk* of AIDS Associated with Blood Transfusion, by Racial/Ethnic Group and Exposure Category

Exposure Category	White*	Black	Hispanic	Other
Transfusion before AIDS diagnosis:				
Men	1.0	1.2 (0.1, 1.7)	1.1 (0.7, 1.6)	0.9 (0.4, 1.7)
Women	1.0	1.9 (1.4, 2.7)	1.4 (0.8, 2.3)	1.2 (0.5, 2.9)
Children	1.0	2.5 (1.4, 4.5)	2.5 (1.2, 5.0)	0.0 (0.0, 2.4)
Heterosexual sex partner had transfusion:				
Men	1.0	24.0 (1.2, 468.0)	0.0 (0.0, 78.1)	0.0 (0.0, 194.9)
Women	1.0	0.5 (0.0, 7.6)	2.1 (0.3, 14.8)	0.0 (0.0, 14.4)
Children whose mother had transfusion	1.0	10.1 (2.1, 48.6)	4.0 (0.4, 37.4)	0.0 (0.0, 34.1)

*Reference group: non-Hispanic Whites (99% confidence interval around relative risk is in parentheses).

TABLE 5—Relative Risk* of AIDS Associated with Undetermined Means of Acquiring HIV Infection (UMAHI), by Racial/Ethnic Group and Exposure Category

Exposure Category	White*	Black	Hispanic	Other
Patients with UMAHI				
Men	1.0	7.2 (6.1, 8.5)	6.3 (5.2, 7.8)	1.2 (0.6, 2.2)
Women	1.0	12.0 (8.7, 16.6)	8.5 (5.5, 13.0)	0.7 (0.1, 4.5)
Children	1.0	7.8 (2.9, 21.0)	5.1 (1.5, 17.8)	0.0 (0.0, 12.4)
Patients with no risk factor for AIDS other than a sex partner with HIV infection but UMAHI				
Men	1.0	9.0 (3.6, 22.5)	1.9 (0.3, 12.9)	0.0 (0.0, 13.0)
Women	1.0	8.6 (4.5, 16.3)	3.4 (1.2, 10.1)	2.4 (0.4, 15.7)
Children with no risk factor for AIDS other than a mother with HIV infection but UMAHI	1.0	24.1 (6.8, 85.6)	3.2 (0.4, 27.7)	4.9 (0.3, 81.7)
Children with no risk factor for AIDS other than a mother whose sex partner had HIV infection but UMAHI	1.0	10.1 (2.1, 48.6)	10.1 (1.8, 56.5)	6.1 (0.3, 108.3)

*Reference group: non-Hispanic Whites (99% confidence interval around relative risk is in parentheses).

from greater use of "shooting galleries" by Blacks and Hispanics, where needles and syringes may be shared with a greater number of persons.*

Blacks' and Hispanics' higher cumulative incidence of AIDS associated with male homosexual activity are not easily explained, as no data are available on the prevalence of homosexuality by race/ethnicity. A study of homosexual men in San Francisco found that Blacks had a higher prevalence and incidence of HIV infection than Whites, which could not be explained by differences in number of sex partners, frequency of receptive anal/genital contact, or sharing needles for drug abuse.[24] Perhaps the higher prevalence of HIV infection in Black and Hispanic IVDAs, noted above, has led to a higher infection rate in their sex partners, whether homosexual or heterosexual.

*Personal communication: L. R. Petersen.

Blacks' and Hispanics' higher cumulative incidence of AIDS associated with male bisexuality could be due to a greater proportion of bisexuals among Black and Hispanic homosexual/bisexual men than among White homosexual/bisexual men.[25] This would be consistent with our finding of a higher risk of AIDS in Black and Hispanic women whose sex partner was a bisexual man, but is not yet significantly reflected in Black and Hispanic children whose mothers' sex partners were bisexual.

Blacks' and Hispanics' higher risk of AIDS in children who have received blood transfusions is probably due to their having a rate of low birth weight at least twice that of Whites.[26] Such births are associated with a higher rate of transfusion for treatment of iatrogenic anemia due to blood sampling in neonatal intensive care units.[27] Blacks' higher risk of AIDS in transfused women, their sex partners, and their children (Table 4) might be explained if Black women

TABLE 6—Cumulative Incidence* and Relative Risk† of AIDS, by Racial/Ethnic Group, Exposure Category, and Geographic Region

Exposure Category	US Region	White		Black		Hispanic		Other	
		CI	RR	CI	RR	CI	RR	CI	RR
Exclusively homosexual men without IVDA‡	Northeast	376.8	1.0	893.2	2.4 (2.2, 2.6)	1,242.2	3.3 (3.0,3.6)	152.5	0.4 (0.3, 0.6)
	Midwest	88.9	1.0	217.7	2.4 (2.1, 2.8)	180.6	2.0 (1.5, 2.7)	34.6	0.4 (0.2, 0.9)
	South	255.2	1.0	248.8	1.0 (0.9, 1.1)	390.4	1.5 (1.4, 1.7)	35.2	0.1 (0.1, 0.3)
	West	657.5	1.0	747.2	1.1 (1.0, 1.3)	400.8	0.6 (0.6, 0.7)	126.3	0.2 (0.2, 0.2)
Bisexual men with IVDA	Northeast	49.3	1.0	341.4	6.9 (6.0, 8.0)	283.0	5.7 (4.7, 7.0)	37.3	0.8 (0.3, 1.7)
	Midwest	21.5	1.0	114.1	5.3 (4.3, 6.7)	57.9	2.7 (1.6, 4.6)	11.5	0.5 (0.1, 2.4)
	South	43.4	1.0	115.9	2.7 (2.3, 3.1)	65.4	1.5 (1.1, 2.0)	8.8	0.2 (0.0, 0.9)
	West	86.6	1.0	245.3	2.8 (2.3, 3.5)	100.4	1.2 (1.0, 1.4)	33.7	0.4 (0.3, 0.6)
Heterosexual adults with IVDA	Northeast	36.8	1.0	951.3	25.9 (23.7, 28.2)	1,128.8	30.7 (27.9, 33.7)	15.0	0.4 (0.2, 1.0)
	Midwest	2.1	1.0	43.5	20.6 (14.6, 29.0)	47.8	22.6 (13.8, 37.0)	3.8	1.8 (0.3, 11.2)
	South	6.0	1.0	92.4	15.4 (12.9, 18.5)	20.0	3.4 (2.3, 4.8)	2.8	0.5 (0.1, 3.0)
	West	7.3	1.0	82.5	11.3 (8.5, 15.2)	23.6	3.2 (2.4, 4.5)	2.3	0.3 (0.1, 1.0)
Adults with undetermined means of acquiring HIV infection	Northeast	5.5	1.0	78.7	14.3 (11.2, 18.3)	91.6	16.7 (12.6, 22.0)	13.3	2.4 (1.0, 6.1)
	Midwest	1.9	1.0	9.1	4.9 (2.9, 8.3)	10.8	5.8 (2.3, 14.3)	1.9	1.0 (0.1, 13.4)
	South	4.5	1.0	24.7	5.5 (4.3, 7.1)	15.8	3.5 (2.4, 5.3)	7.1	1.6 (0.5, 5.1)
	West	6.0	1.0	45.5	7.6 (5.3, 10.9)	15.2	2.5 (1.7, 3.7)	2.3	0.4 (0.1, 1.2)
All AIDS patients (including children)	Northeast	218.4	1.0	1,445.9	6.6 (6.3, 6.9)	1,538.8	7.0 (6.7, 7.4)	100.0	0.5 (0.3, 0.6)
	Midwest	53.1	1.0	171.8	3.2 (2.9, 3.6)	138.6	2.6 (2.1, 3.2)	28.1	0.5 (0.3, 0.9)
	South	143.9	1.0	286.7	2.0 (1.9, 2.1)	206.3	1.4 (1.3, 1.6)	31.4	0.2 (0.1, 0.3)
	West	351.0	1.0	560.9	1.6 (1.5, 1.7)	227.5	0.6 (0.6, 0.7)	71.5	0.2 (0.2, 0.2)

*Cumulative incidence (CI): AIDS cases reported from June 1, 1981, to January 18, 1988, per million population.
†Reference group for relative risk (RR): non-Hispanic Whites (99% confidence interval around RR is in parentheses).
‡History of intravenous-drug abuse.

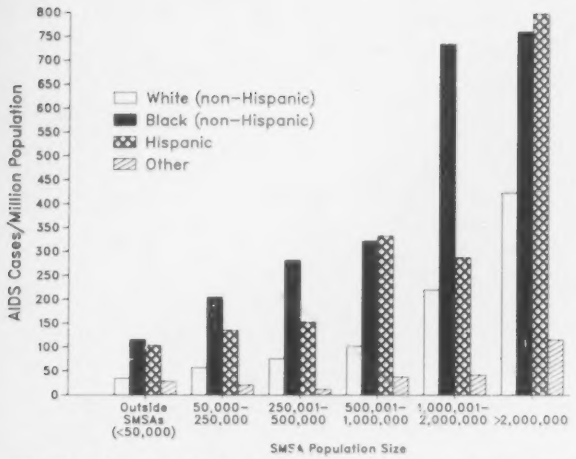

FIGURE 1—Cumulative Incidence of AIDS as of January 18, 1988 by Racial/Ethnic Group and Population Category of Standard Metropolitan Statistical Areas (SMSAs), based on the 1980 US Census

had a higher incidence of transfusions than White women. If they do, their higher transfusion rate might be due to their higher rates of birth, abortions, and ectopic pregnancies.[28–30] Alternatively, since HIV transmission by transfusion is generally only presumed in these cases, without being confirmed by assessing the HIV infection status of the blood donors, the history of blood transfusion may be merely a coincidental finding in a proportion of these cases.

Black men's lower risk of AIDS related to hemophilia suggests that this inherited disorder may be less common in Blacks than Whites, but demographic data on hemophiliacs are unavailable to substantiate this.**

Blacks' and Hispanics' higher risk of AIDS associated with undetermined means of acquiring HIV infection could result if Blacks and Hispanics were more reluctant to admit to IVDA or homosexual activity than Whites. Even without a higher rate of concealment of risk factors, however, a disproportionately high number of Black and Hispanic AIDS patients in the undetermined categories should be expected simply due to their higher proportions in the IVDA and homosexual/bisexual categories. In addition to concealing IVDA or homosexual activity in AIDS patients, these undetermined categories might reflect heterosexual transmission to persons unaware that their sex partners were IVDAs or bisexual men or had other sex partners with risk factors. Such heterosexual transmission might be occurring at a greater rate in Blacks and Hispanics than Whites.

The focus of this study has been on RRs, because they reveal differences not readily apparent from the percentage distribution of AIDS cases by race/ethnicity and exposure category. However, the importance of racial/ethnic differences in risks is reinforced by the magnitude of the proportion of Blacks and Hispanics in certain categories of AIDS patients. Of AIDS cases associated with IVDA by heterosexuals (including sex partners and children of IVDAs), more than 80 per cent were in Blacks and Hispanics. The preponderance of Blacks and Hispanics among heroin addicts was noted more than 20 years ago.[31] In addition, Blacks and Hispanics constitute a disproportionate 48 per cent of AIDS cases in women whose male sex partners were bisexual and 59 per cent of cases with undetermined means of acquiring HIV infection.

The prevalence of high-risk behavior may differ between persons in different areas, despite their belonging to the same racial/ethnic group as broadly categorized here. For example, if high-risk homosexual activity were more prevalent in Whites or less prevalent in Hispanics in the western US than in their counterparts in other regions, either situation could explain why the RR in Hispanics is lower in the west than in other regions (Table 6). The same principle could explain why, in the central cities, the general populations of Blacks and Hispanics have a lower cumulative incidence than Whites of AIDS in exclusively homosexual men (Table 7). Since our comparison of central cities with suburbs was limited to SMSAs with ≥2,000,000 population, the differences observed may not necessarily reflect the situation in smaller SMSAs or outside SMSAs. Independent of the prevalence of high-risk behavior, the prevalence of HIV infection may differ between regions due to differences in the time when the epidemic began in particular exposure categories.

Our analysis of trends was too early to assess the impact of the 1987 revision of the AIDS case definition on relative risks, as reporting of cases under the revised definition did not begin until September 1, 1987.[10]

The possibility that the observed higher risks in Blacks and Hispanics may be due to biased data needs to be considered. The proportion of the population not counted in the 1980 census was greater for Blacks†[32] and Hispanics‡ than Whites. When we modified our calculations, however, by increasing the populations by estimates of their undercounts, RRs in Blacks and Hispanics were not significantly reduced. If private physicians were more concerned than other physicians about guarding the confidentiality of their patients, reporting of AIDS cases might be more complete in indigent (disproportionately Black and Hispanic) patients of public hospitals than in patients who can afford private medical care. A bias in the opposite direction, however, may result if indigent patients were less likely to undergo expensive diagnostic procedures needed to meet the AIDS case definition (presumptive diagnoses were not accepted until September 1987).[10] A reporting bias would probably be too small to explain the magnitude of the observed RRs, since a validation study estimated the completeness of AIDS reporting to be about 90 per cent, a rate that did not vary by racial/ethnic group.[33] Furthermore, a reporting bias would be unlikely to explain fully the variation in RRs by exposure category and geographic area.

HIV-antibody tests in applicants for military service[34] and in volunteer blood donors[35] have shown that the prevalence of HIV infection is higher in Blacks and Hispanics than in Whites, but not higher in American Indians/Alaskan Natives and Asians/Pacific islanders. This supports our analyses. Awareness that the risk of AIDS is not higher in racial/ethnic minorities other than Blacks and Hispanics

**Literature search by R. M. Selik and personal communication from J. Jason.

†Passel JS, Robinson JG: Bureau of the Census memorandum, April 8, 1985.

‡Passel JS, Woodrow KA: Bureau of the Census memorandum, January 16, 1987.

TABLE 7—Cumulative Incidence* and Relative Risk† of AIDS, by Racial Ethnic Group, Exposure Category, and Residence in the Central Cities or Suburbs of Standard Metropolitan Statistical Areas of ≥2 Million Population

Exposure Category	Residence	Racial/Ethnic Group	AIDS Cases	Cumulative Incidence	RR	(99% Confidence Interval)
Exclusively homosexual men	Central cities	White	12,192	1,013.4	1.0	
		Black	2,638	346.8	0.3	(0.3, 0.4)
		Hispanic	1,760	511.1	0.5	(0.5, 0.5)
		Other	128	125.9	0.1	(0.1, 0.2)
	Suburbs	White	2,686	88.1	1.0	
		Black	387	140.3	1.6	(1.4, 1.8)
		Hispanic	261	113.3	1.3	(1.1, 1.5)
		Other	23	23.3	0.3	(0.2, 0.5)
Bisexual men	Central cities	White	1,128	93.8	1.0	
		Black	1,050	138.0	1.5	(1.3, 1.6)
		Hispanic	378	109.8	1.2	(1.0, 1.4)
		Other	25	24.6	0.3	(0.2, 0.4)
	Suburbs	White	588	19.3	1.0	
		Black	169	61.3	3.2	(2.5, 4.0)
		Hispanic	86	37.3	1.9	(1.4, 2.6)
		Other	9	9.1	0.5	(0.2, 1.1)
Heterosexual adults with IVDA‡	Central cities	White	608	50.5	1.0	
		Black	2,401	315.6	6.2	(5.6, 7.0)
		Hispanic	1,537	446.4	8.8	(7.8, 10.0)
		Other	8	7.9	0.2	(0.1, 0.4)
	Suburbs	White	258	8.5	1.0	
		Black	491	178.0	21.0	(17.3, 25.6)
		Hispanic	133	57.7	6.8	(5.2, 9.0)
		Other	5	5.1	0.6	(0.2, 1.9)
Adults with undetermined means of acquiring HIV infection	Central cities	White	157	13.0	1.0	
		Black	298	39.2	3.0	(2.3, 3.9)
		Hispanic	166	48.2	3.7	(2.8, 4.9)
		Other	9	8.9	0.7	(0.3, 1.6)
	Suburbs	White	119	3.9	1.0	
		Black	68	24.7	6.3	(4.3, 9.3)
		Hispanic	40	17.4	4.4	(2.8, 7.1)
		Other	4	4.0	1.0	(0.3, 3.8)
All AIDS patients (including children)	Central cities	White	14,379	1,195.2	1.0	
		Black	7,246	952.5	0.8	(0.8, 0.8)
		Hispanic	4,126	1,198.2	1.0	(1.0, 1.0)
		Other	187	183.9	0.2	(0.1, 0.2)
	Suburbs	White	4,038	132.5	1.0	
		Black	1,306	473.5	3.6	(3.3, 3.9)
		Hispanic	571	247.8	1.9	(1.7, 2.1)
		Other	49	49.6	0.4	(0.3, 0.5)

*Cumulative incidence = AIDS cases reported from June 1, 1981 to January 18, 1988 per million population of the same race/ethnicity.
†Reference group for relative risk: non-Hispanic Whites.
‡History of intravenous-drug abuse.

should help stop misleading generalizations that imply higher risks in all "minority" groups.[2,3,18]

Black and Hispanic communities should be targetted for special efforts to prevent and treat IVDA and to counsel IVDAs on the risk of HIV infection. The need for this is greatest in the northeast. Recommendations for preventing HIV transmission to IVDAs, their sex partners, and their children have been published.[36-41] Black and Hispanic women need to know that they can acquire HIV infection by sexual contact with male IVDAs. Prevention messages to Black and Hispanic men must be addressed not only to men who identify themselves as gay, but also to men who see themselves as basically heterosexual despite occasional sexual relations with other men. For these and other recommendations to be implemented, they must be endorsed by Black and Hispanic community leaders and others whose judgment is respected by the persons for whom the messages are intended.

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New Quarterly *Journal of Human Lactation* Announced

A new quarterly journal—the *Journal of Human Lactation*—has been launched by the International Lactation Consultant Association (ILCA) to provide current information on the latest clinical concerns relating to human lactation and breastfeeding. Features include research and scientific articles, commentaries focusing on controversial issues, and brief reports highlighting the clinical application of new knowledge about lactation and the breastfeeding encounter. In addition, *JHL* includes listings of relevant articles from other sources, offers discussions of practice-oriented concerns of lactation consultants in a variety of settings, and news of political/governmental/marketing activities that influence breastfeeding women worldwide.

Kathleen G. Auerbach, PhD, is Editor-in-Chief of *JHL*. The editorial offices are located at 2240 Willow Road, Homewood, IL 60430-3221. Subscriptions ($40 individual; $50 institutions) are handled by ILCA, Department PH, PO Box 4031, University of Virginia Station, Charlottesville, VA 22903 USA.

Prevalence of Cardiovascular Disease Risk Factors in Blacks and Whites: The Minnesota Heart Survey

J. Michael Sprafka, PhD, Aaron R. Folsom, MD, Gregory L. Burke, MD, and Stanley A. Edlavitch, PhD

Abstract: Two cross-sectional surveys were conducted in 1985 and 1986 to measure the prevalence of coronary heart disease (CHD) risk factors in Blacks and Whites. A home interview was followed by a survey center visit. Participation rates were 78 per cent and 90 per cent for the home interview and 65 per cent and 68 per cent for the survey center visit. Adjusted for age and education, systolic and diastolic blood pressure was 3 to 4 mmHg higher in Blacks. Hypertension was more prevalent in Blacks than Whites (44 per cent vs 28 per cent); serum total cholesterol was approximately 0.4 mmol/l lower in Black than White men and 0.08 mmol/l lower in Black than White women. Among men, more Blacks than Whites were current cigarette smokers (44 per cent vs 30 per cent); however, White smokers smoked more cigarettes per day (26 vs 17). Similar differences were noted for women, although the prevalence and quantity of cigarette consumption was less than men. The excess prevalence of these CHD risk factors in Blacks, especially among women, may explain their elevated CHD and stroke mortality rates in the Twin Cities. (*Am J Public Health* 1988; 78:1546–1549.)

Introduction

The relation of blood pressure, serum cholesterol, and smoking to coronary heart disease (CHD) incidence and mortality rates has been reported in many epidemiologic studies both nationally and internationally.[1–3] CHD mortality rates are higher for Blacks than Whites at younger ages, and lower for Blacks after the age of 70.[4–6] Whether racial differences reflect access to medical care or prevalence of CHD risk factors is unclear. Few reports compare the prevalence of the known CHD risk factors between Blacks and Whites. This paper describes the prevalence of coronary heart disease risk factors in Blacks and Whites using data from two population-based surveys in a large metropolitan area.

Methods

Study Population

Methods and early findings from the Minnesota Heart Survey, a survey of CHD risk factors and related behaviors in probability samples of Twin Cities residents aged 25 to 74 years, have been described previously.[7] The seven-county metropolitan area was divided into 704 clusters of approximately 1,000 households each, based on 1970 and 1980 census information, from which 40 clusters and approximately 10 per cent of households within each cluster were randomly selected. One individual aged 25–74 was randomly selected per household, interviewed in the home, and then invited to attend a neighborhood clinic for more extensive interviews and physiologic tests. The first series of general population surveys was conducted in 1980–82. A second series was begun in December 1985, with two complete one-year survey cycles.

From July to November 1985, we also surveyed an area of Minneapolis with large numbers of Black residents to provide more precise estimates of their CHD risk factors. In the sampling area, which included four and one-half census tracts, every occupied housing unit was identified and *all* 35- to 74-year-old Blacks were considered eligible and recruited to the survey. Other methods were identical to those in the general population survey; preliminary findings are detailed elsewhere.[8,9]

Data on Risk Factors

In both surveys, height was measured in stocking feet to the nearest half inch. Body mass index (BMI) was computed as $wt/ht^2 (kg/m^2)$. Blood pressure was measured by trained technicians using a random zero sphygmomanometer on the right arm of seated participants after a five-minute rest. The average of two measures of systolic and diastolic (5th phase) is reported. Non-fasting blood specimens were obtained using the protocol of the Lipid Research Clinics.[10] The Jenkins Activity Survey[11] was used to assess Type A behavior (positive scores indicate more Type A tendencies and negative scores more Type B). The Minnesota Leisure Time Physical Activity (LTPA) questionnaire[12] was administered to a systematic 50 per cent sample of study participants. The systematic 50 per cent not receiving the LTPA were given a 24-hour dietary recall questionnaire. Exertional anginal pectoris,[13] smoking, alcohol use, demographics, and medication use were assessed by questionnaire.

Analysis

Data on Whites aged 35 to 74 from the first cycle of the 1985–87 general population survey were compared to those obtained from the 1985 Black population survey. We used analysis of variance to adjust for the effects of age and education to determine the relationship and overall significance of racial differences.[14]

Results

Participation rates in the general population survey were 90 per cent for the home interview alone and 68 per cent for the home interview and clinic visit. Participation rates in the 1985 Black survey were 75 per cent and 65 per cent, respectively. In both surveys individuals attending clinic were more likely than non-clinic attenders to be non-smokers (47.7 per cent vs 63.2 per cent), have more education (greater than high school education, 54.0 per cent vs 48.1 per cent), be married (63.8 per cent vs 57.9 per cent) and be currently employed (71.1 per cent vs 66.3 per cent). No differences were noted in the percentages of individuals with a history of high blood pressure, hypercholesterolemia, diabetes, heart attack, or stroke or in leisure time physical activity scores between clinic attenders and non-attenders. Table 1 shows

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TABLE 1—Sociodemographic Characteristics of Men and Women Ages 35–74 Years by Race Adjusted for Age: The Minnesota Heart Survey

Variables	Men		Women	
	White (N=893)	Black (N=547)	White (N=976)	Black (N=706)
Age (years)	49.1	50.1	49.9	50.8
% Married	78.0	63.4	62.9	43.2
% Currently employed	81.3	71.4	64.8	63.5
Educational Status				
% ≤ high school	34.9	54.2	46.9	56.8
% some college	30.3	28.5	30.3	29.1
% ≥ college degree	34.9	17.3	22.8	14.1
Income Status†				
% < $15,000	6.7	25.5	17.0	40.2
% $15,000–$30,000	21.0	34.8	29.5	32.2
% > $30,000	70.6	38.3	49.9	25.0
% Unavailable	1.7	1.3	3.6	2.6

†For income status N = 763–WM; 457–BM; 811–WW; 590–BW.

the sociodemographic characteristics of men and women by race adjusted for age.

The age-adjusted prevalence of diabetes, angina, history of heart attack or stroke, and average body mass index was higher in Black than White women (Table 2). No racial difference in these conditions was noted for men. In general, scores for physical activity and Type A behavior were higher in Whites than Blacks. Alcohol consumption was higher in Black than White men with no differences noted for women.

For all age groups, systolic and diastolic blood pressures were higher in Blacks than Whites by 3 or 4 mmHg (Table 3). These differences remained after stratification by antihypertensive medication status. The prevalence of hypertension, defined as systolic blood pressure greater than or equal to 140 mmHg and/or diastolic blood pressure greater than or equal to 90 mmHg and/or currently taking antihypertensive medication, was higher in Blacks than Whites for all age-sex categories (Table 4).

Figure 1 shows that a greater percentage of Black than White men were aware of their hypertension, treated and controlled. More Black than White men had treated but uncontrolled hypertension, except 65- to 74-year-olds. The percentage of men aware of their hypertension yet untreated and uncontrolled was similar in Blacks and Whites and decreased consistently with age. Overall, the percentage of men unaware of their hypertension was markedly higher among Whites than Blacks (38 per cent vs 27 per cent: mean difference = 10.5 per cent, 95 per cent CI = 1.7, 19.3).

Compared to men, women were more aware of their hypertension (Figure 2) and it was better treated and con-

TABLE 3—Average Blood Pressure (mm/Hg) for Men and Women by Race and Age: The Minnesota Heart Survey

Age (years)	Men		Women	
	White	Black	White	Black
Systolic				
35–44	122.6	125.1	113.1	116.2
45–54	125.6	129.2	119.5	124.1
55–64	129.5	133.0	127.9	132.4
65–74	135.1	137.1	133.7	134.9
Total†	126.7	129.0	120.9	124.3
Mean difference† (95% CI)	2.3(0.5,4.3)		3.4(1.6,5.2)	
Diastolic				
35–44	76.7	79.8	71.5	75.5
45–54	79.2	83.5	74.3	78.9
55–64	76.9	79.7	75.4	79.3
65–74	74.9	81.6	72.3	75.5
Total†	78.7	82.4	74.1	78.2
Mean difference† (95% CI)	3.7(2.3,5.1)		4.1(2.9,5.3)	

†Age and Education-adjusted.

TABLE 4—Prevalence of Hypertension† among Men and Women by Age and Race: The Minnesota Heart Survey

Age (years)	Men		Women	
	White	Black	White	Black
Systolic				
35–44	19.3	30.4	7.8	21.9
45–54	28.4	39.6	21.9	42.9
55–64	42.7	57.0	40.0	68.6
65–74	47.0	70.3	57.8	68.7
Total*	30.7	43.7	24.8	44.1
Mean difference* (95% CI)	12.9(7.5,18.3)		18.6(14.1,23.1)	

†Hypertension defined as systolic blood pressure ≥ 140 and or diastolic blood pressure ≥ 90 mm Hg and or currently taking antihypertensive medication.
*Age and Education-adjusted.

TABLE 2—Medical Conditions and Behavior Patterns by Race and Sex Adjusted for Age: The Minnesota Heart Survey

Variables	Men			Women		
	Black	White	95% CI of Difference	Black	White	95% CI of Difference
History of diabetes (%)	7.3	4.8	(−0.002,4.6)	12.0	4.2	(5.4,10.4)
History of angina (%)	2.7	3.2	(−0.01,0.02)	6.8	3.5	(1.2,5.4)
Prior hospitalization for heart attack or stroke (%)	6.0	5.2	(−0.02,0.03)	5.1	2.4	(0.9,4.5)
Body mass index (Kg/m²)	27.3	27.3	(−0.5,0.5)	29.0	26.2	(2.2,3.4)
LTPA*	2.1	2.3	(0.1,0.3)	1.9	2.1	(0.06,0.20)
JAS**						
Working	−4.8	−2.3	(1.3,3.6)	−5.1	−5.0	(−0.9,1.3)
Non-working	−9.2	−8.3	(−1.9,3.7)	−7.0	−8.4	(0.3,3.1)
Alcohol consumption (oz/wk)	7.3	4.6	(1.7,3.7)	1.9	1.9	(−0.4,0.5)

*Log transformed Leisure Time Physical Activity Score.[12]
**Jenkins Activity Survey Score.[11]

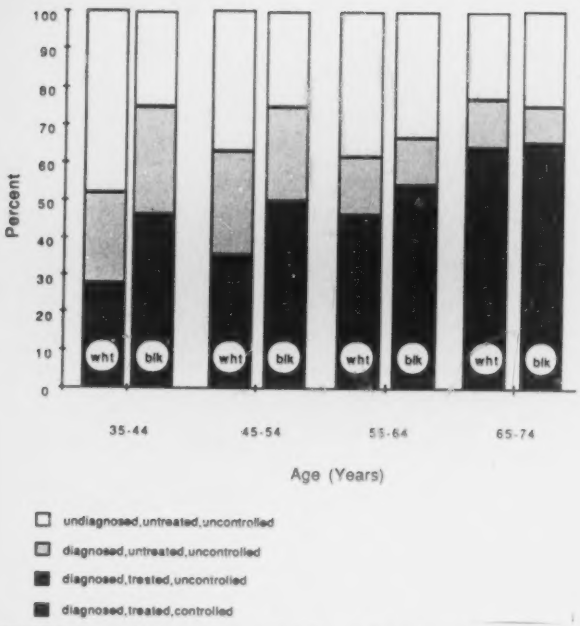

FIGURE 1—Awareness, Treatment and Control Status among Hypertensive Men by Race and Age: The Minnesota Heart Survey

trolled in all ages. The fraction of women diagnosed, treated and uncontrolled averaged approximately 25 per cent in both Blacks and Whites. More White than Black women were unaware of their hypertension (24 per cent vs 15 per cent: mean difference = 8.9 per cent, 95 per cent CI = 1.7, 16.1).

In general, Black men had lower serum total cholesterol and higher serum HDL cholesterol levels than White men

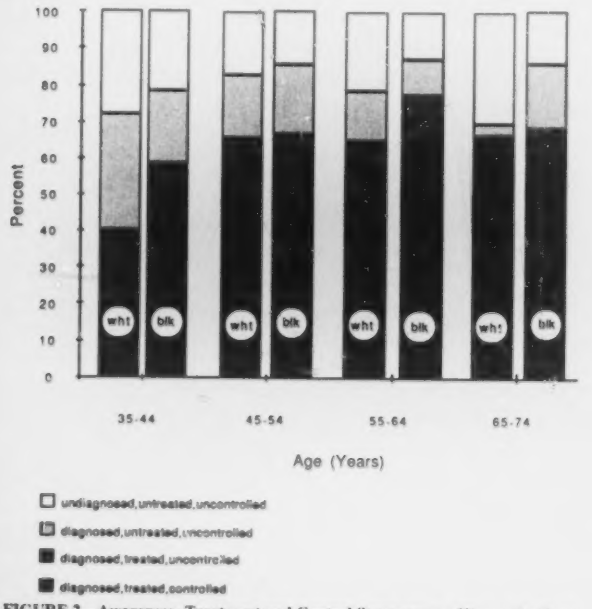

FIGURE 2—Awareness, Treatment and Control Status among Women by Race and Age: The Minnesota Heart Survey

TABLE 5—Serum Total and HDL Cholesterol Values (mg/dl)* for Men and Women by Age and Race: The Minnesota Heart Survey

Age (years)	Total Cholesterol (TC)		HDL Cholesterol		HDL/TC Ratio	
	Black	White	Black	White	Black	White
Men						
35–44	186.7	203.0	48.2	40.0	0.26	0.20
45–54	196.2	209.7	47.8	42.2	0.25	0.21
55–64	191.7	212.1	49.1	40.0	0.26	0.20
65–74	207.6	201.5	50.6	42.4	0.25	0.22
Total†	193.2	206.8	48.6	40.9	0.26	0.21
Mean Difference† (95% CI)	13.6(9.0,18.2)		7.7(6.1,9.3)		0.05(0.04,0.06)	
Women						
35–44	184.2	185.8	55.8	54.8	0.31	0.30
45–54	201.2	203.5	54.7	54.8	0.28	0.28
55–64	218.0	226.0	58.3	53.2	0.27	0.24
65–74	222.7	235.6	55.8	50.9	0.25	0.23
Total†	201.4	206.0	56.3	53.8	0.29	0.27
Mean Difference† (95% CI)	4.6(2.5,6.7)		2.5(0.9,4.1)		0.02(0.003,0.02)	

†Age and Education-adjusted.
*1 mg/dl = 0.02586 mmol/l.

(Table 5). Black women had slightly lower serum total cholesterol values than Whites among individuals aged 35 to 54 and the differences increased with age. Black women had slightly higher values of serum HDL cholesterol than White women in the younger age groups and this difference increased with age. The ratio of HDL to total serum cholesterol remained relatively stable with age in men but gradually declined in women.

The age- and education-adjusted prevalence of cigarette smoking was higher in Black than White men (43.9 per cent vs 29.5 per cent) and women (27.9 per cent vs 17.9 per cent).* Black and White men were equally likely ever to have smoked, but many more of the White men had quit. More Black than White women were current cigarette smokers especially among those aged 35 to 44. Black women were more likely never to have smoked than White women, but fewer Black women had quit.

Discussion

The effect of non-response on the CHD profiles in this study is unknown; however, previous reports indicate that survey non-participants more often are unmarried, unemployed, smokers, and have alcohol problems than participants.[7,15] Our findings on clinic participation are consistent with these observations. Overall response rates were similar in the Black and general population surveys, 65 per cent and 68 per cent, respectively.

With the exception of serum total and HDL cholesterol levels and Jenkins Activity Survey scores,[11] Blacks had a relative excess of the known CHD risk factors. The results regarding blood pressure, blood lipids, and cigarette consumption differences between Blacks and Whites agree closely with those reported elsewhere.[16–19] If, as our data suggest, these blood pressure differences between Blacks and Whites are not attributable to lower hypertension awareness and treatment patterns among Blacks, one possible explanation is that Blacks have more severe hypertension with earlier

*Data available on request to author

onset and poorer control when treated. Differences between Blacks and Whites in hypertension awareness and control in this survey are consistent with the National Health and Nutrition Examination Survey (NHANES) data[20]; but the percentages of individuals aware, treated and controlled are much higher than those reported in NHANES for both men and women, Black and White, suggesting that Minnesota may have more effective hypertension control programs than the US as a whole.

Average serum cholesterol levels among US adults have declined significantly in the last 20 years for Whites but not for Blacks.[21] The average serum cholesterol level in Blacks is lower than in Whites in Minnesota and nationwide, yet Blacks in this study had higher serum HDL cholesterol values than Whites. Reasons for this are unclear and are not likely explained by differences in alcohol and cigarette consumption, obesity, or physical activity. A genetic component, as suggested by Gartside, *et al*,[23] might be an hypothesis needing further exploration.

The excess of CHD risk factors among Twin Cities Blacks, especially women, may partially explain their elevated cardiovascular mortality rates.[8] Despite significant declines during the last decade for all sex- and race-specific groups, heart disease mortality rates are higher in Blacks than Whites (10 per cent excess in men, 60 per cent excess in women). In addition, stroke mortality rates are higher in Blacks than Whites (23 per cent excess in men; 79 per cent excess in women). Public health interventions must focus on reducing these risk factors if further improvements in Black CHD mortality rates are to be seen.

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Behavioral Aspects of Cigarette Smoking among Industrial College Men of Shanghai, China

VIRGINIA C. LI, PHD, HU JIN-HUA, ZHOU MANLAN, AND ZHENG JIABAO

Abstract: A questionnaire survey was conducted among 7,665 students and staff in the factory-operated industrial colleges in Shanghai, China. Response rate was 95.7 per cent. The proportion of smokers among the men was 50.5 per cent; only 0.33 per cent of the women reported they were smokers. Male smokers began smoking upon entering the work force, where social interaction encourages cigarette smoking. Seventy per cent of the male smokers smoked fewer than 10 cigarettes per day. Only 4.3 per cent perceived themselves as very addicted; 3.4 per cent reported they had no confidence in their ability to control the amount they smoke. The Chinese smokers are more likely to succeed in giving up cigarette smoking for at least one day, as well as for longer periods, compared to US smokers. The number of cigarettes smoked may be related to the individual's purchasing power. For China, strategic timeliness in legislative action and education is urgently needed to combat the high prevalence of cigarette smoking in males, at extremely high risk. (*Am J Public Health* 1988; 78:1550–1552.)

Introduction

China leads the world in tobacco production. Cigarette manufacturing and sales are state monopolies; annual profit to the state from the tobacco industry constitutes 8 per cent of national income and is one of the largest sources of revenues for the Chinese government.[1] In 1987, of the 50 largest industrial enterprises in China, eight were manufacturers of tobacco products.[2]

There are indications that lung cancer and ischemic heart diseases are on the rise in China.[3,4] In recent years, China's State Patriotic Health Campaign Commission has assumed an advocacy role in smoking prevention.[5] Since the Commission is an arm of the State Council, the highest executive body of the central government, this suggests that the State recognizes the health consequences of smoking. Efforts to call public attention to the health risks of smoking are increasing.[6]

The World Health Organization-Shanghai Collaborating Center for Health Education was established in 1985 to strengthen health education and the delivery of primary care to local communities. The Center supported a systematic survey to determine smoking prevalence, knowledge, attitudes, and behavioral characteristics of the students and staff of factory-operated industrial colleges in the municipality of Shanghai. This paper reports the survey findings.

Methods

Study Population

Twenty-six of the 54 factory-operated industrial colleges in Shanghai participated in the study, with a total of 7,665 students and staff. The industrial colleges are three-year technical schools offering specialization in various occupations. They have from 107 to 671 students, and a student/staff ratio of 2:1. The students are older than general university students. They have had at least one year of practical work experience.

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Procedures

A self-administered structured questionnaire was designed to determine demographic characteristics, cigarette smoking knowledge, attitudes and behavior, and reasons for cigarette smoking, and smoking and quitting. The survey was implemented in January 1986.

Cooperation from the project sites was obtained by the Center staff. The clinic physician or health worker was usually assigned as the site coordinator at each of the participating colleges. Coordinators were trained by the investigators to administer the questionnaire in group settings. When respondents were illiterate (less than 0.3 per cent), coordinators helped them complete the questionnaire.

Results

Of the 7,339 respondents (95.7 per cent of those eligible), 4,899 were men and 2,440 women. The mean age was 30.8 ± 9.7 years for men and 28.7 ± 7.7 years for women.

Chinese smokers are almost exclusively male. The proportion of current smokers among the men was 50.05 per cent (N = 2,447), while only 0.33 per cent (N = 9) of the women reported that they smoked. Among the men, 90 per cent of the non-smokers versus 84 per cent of the smokers had some college level education; marital status of smokers and non-smokers did not differ appreciably. Nearly 90 per cent of the respondents reported they were very satisfied with their work, irrespective of smoking status. Since there were so few female smokers, the remainder of this analysis will focus on the male smokers.

Smoking Behavior

The bulk of the smoking men were light smokers (Figure 1). Of the smokers who reported the amount they smoked (88 per cent), 70 per cent reported that they smoked an average of one to 10 cigarettes per day, and less than 5 per cent smoked one or more packs per day. Smokers used both filter and non-filter cigarettes; 16 per cent of respondents said that they did not inhale, 28 per cent lightly inhaled, 48 per cent usually inhaled, and 8 per cent deeply inhaled.

The onset of cigarette smoking occurred in early adulthood when the individual entered the job market and began to earn an income. Twenty-five per cent of the male smokers started smoking cigarettes between the ages 15–19, 55 per cent between ages 20–24.

Figure 2 shows that older smokers tended to smoke more cigarettes than the younger smokers. The proportion doubled

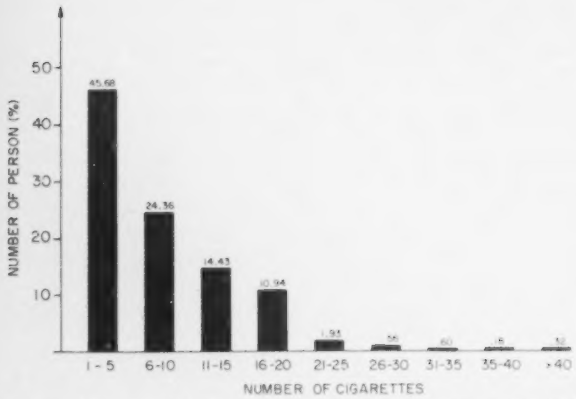


FIGURE 1—Average Number of Cigarettes Smoked among Men in Shanghai Industrial Colleges, 1986

for 11–20 cigarettes, quadrupled for 21–30 cigarettes, and increased ten-fold for 31–40 cigarettes.

Nearly all (96 per cent) of the male smokers were able to identify where they smoked the first cigarette of the day: the workplace (34 per cent), with friends (32 per cent), at home (14.4 per cent), and at school (12.8 per cent). The social aspect of cigarette smoking is even more evident when "friends" are combined with "home of relatives and friends" (35.8 per cent). Only 6 per cent of smokers reported no smokers among their five closest associates, compared to 16.3 per cent of the non-smoking respondents. The male smokers also reported that 63 per cent of their superiors smoked compared to 48 per cent of the non-smoking respondents. With respect to family members who smoke, the proportions of fathers (46.8 per cent), brothers (35.6 per cent), and children (6.1 per cent) of smokers were all higher compared to fathers (39.9 per cent), brothers (24.1 per cent), and children (2.9 per cent) of non-smokers.

Table 1 presents the reasons smokers gave for smoking (93 per cent reported at least one reason). The most common reasons given for smoking were social need (78.7 per cent), problems with work/study (61.3 per cent), and relaxation (54.1 per cent). A relatively low percentage of respondents reported "pleasure" (27.6 per cent) and "desire for cigarettes" (11 per cent). When smokers were asked about their

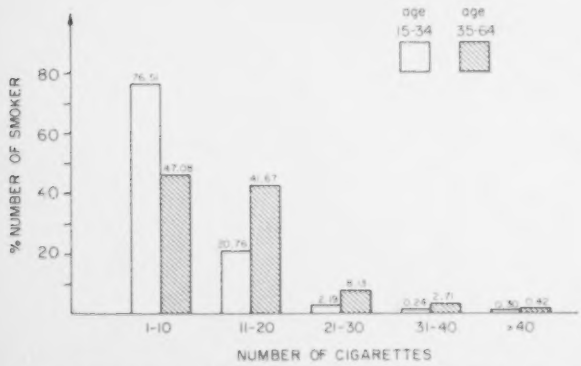


FIGURE 2—Average Number of Cigarettes Smoked per Day by Age, among Men in Shanghai Industrial Colleges, 1986

TABLE 1—Reasons for Smoking and Per Cent Respondents among Men in Industrial Colleges, Shanghai, 1986

Reasons for Smoking	% Respondents	% Respondents Giving Positive Responses
	N = 2,447	
Social needs	80.9*	78.7
Problems with work/study	71.8	61.3
Relaxation	70.4	54.1
Stress Reduction	68.9	51.5
Problems at home	66.8	42.4
Recreation	65.5	24.4
Pleasure	64.8	27.6
Something to do with my hands	63.5	19.4
Desire for cigarettes	62.9	11.0

*Response related to higher educational level for all reasons by χ^2, $p < .005$.

desire to smoke in places where smoking is not appropriate, 96 per cent responded: 65 per cent reported no desire, 31 per cent some desire, and 4 per cent great desire.

Health Knowledge about Cigarette Smoking

Table 2 shows that the majority of both non-smokers and smokers were aware that cigarette smoking may cause cancer and heart disease, and be harmful to the fetus, but more non-smokers than smokers were aware of all these risks.

Over 75 per cent of the non-smokers compared to 37 per cent of the smokers acknowledged that smoking is a distasteful habit. Only about 50 per cent of the respondents, both smokers and non-smokers, felt that cigarette smoking is an addictive habit.

Teachers were consistently more knowledgeable than the other respondents about the hazards of cigarette smoking.

Smoking Cessation

Smokers were asked whether they had been advised by someone to stop cigarette smoking. Of the 93 per cent who responded, 65 per cent had received such advice during the previous year; 98 per cent gave information on whether they had made a serious attempt to quit following the advice, and 39 per cent reported having tried to quit. Figure 3 shows that of those who tried about half were able to quit for six weeks and almost one-third had quit for up to five months before relapsing. In their previous attempts to quit cigarette smoking, 74 per cent said that they did not feel any adverse sensations when quitting, but 26 per cent did experience nervousness and anxiety.

Smokers were queried about the reasons for wanting and not wanting to quit smoking. Fifty-two per cent gave at least one reason for quitting, with "improve health" (66 per cent) being the leading reason. Other reasons were: "too expensive" (44.6 per cent), "bad for health," (42.4 per cent), "not

TABLE 2—Health Hazard Knowledge of Cigarette Smoking among Men in Industrial Colleges, Shanghai, 1986

Health Hazards	Smokers N = 2,262	Non-Smokers N = 5,623
Cigarettes may cause cancer	70.3	86.5
Cigarettes may cause heart disease	55.1	62.3
Passive smoking is harmful to unborn child	75.6	87.8
Passive smoking is harmful to health	72.5	87.5

want to smoke" (36 per cent), "bad habit" (37.7 per cent), "advised to quit" (31.8 per cent), and "bad example" (17.5 per cent).

As to specific methods of quitting smoking, "using will power" was cited by 74 per cent of the respondents. Other methods included "substitution by food and drinks," "reminder by family and friends," "forming a quit smoking group," and "self-monitoring." However, the proportions who reported having used these methods were small, ranging from 18.6 per cent to 9.7 per cent.

Consistent with the data on the number of cigarettes smoked per day, only 4.3 per cent of the smokers perceived that they are very addicted, and 3.4 per cent reported that they have no confidence in their ability to control the amount of cigarettes they smoke. The proportion who felt that they were totally in control of the amount of cigarettes they smoked (56.4 per cent) was greater than those who indicated they were not addicted (45.3 per cent). This suggests that many who consider themselves somewhat addicted are also confident that they are in control of their smoking habits. Eight per cent of the smokers predicted that they would be ex-smokers one year from the time of the study. Nearly all smokers (97 per cent) responded to the question on their desire to smoke in places where smoking is not appropriate. Significantly, 65 per cent said "no", 31 per cent said "sometimes", and 4 per cent said "frequently".

Discussion

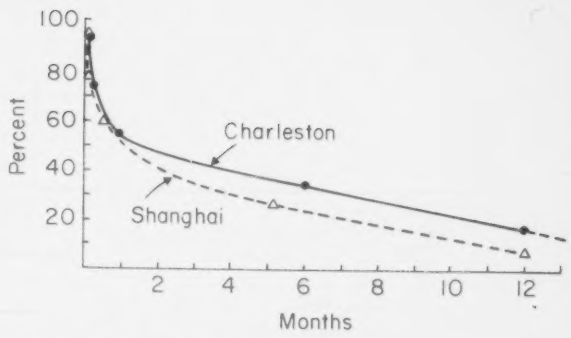
The proportion of smokers reported in the Shanghai industrial colleges is lower than that reported for the general Chinese population,[5] perhaps because of the higher education levels of the Shanghai sample.

The Shanghai data revealed the enormous discrepancy in smoking rates between men and women in China. Why, in a society that is ideologically egalitarian, is there such a marked discrepancy between the sexes in smoking rates, when these rates are increasingly becoming equal between the sexes in western countries?

The strongest influence on women smoking in the west has been commercial advertising. Beginning in the 1920s, tobacco companies in the west have systematically targeted women, capitalizing on their growing income and sense of independence.[9,10] After World War II, women constituted one of the fastest growing segments of the tobacco market.[11] Commercial advertising is restricted and women are largely untouched by commercial exploitation in socialist China. But even in pre-revolutionary China, female cigarette smoking was frowned upon, often associated with prostitutes, and thus traditionally unacceptable.

In western countries those who have not adopted the smoking habit by age 20 are unlikely ever to do so. In Shanghai, however, only 25 per cent of adult smokers had started to smoke before age 20 and none reported having smoked before age 15. There is a strict policy to expel any students from high schools for smoking cigarettes. Also, Chinese children seldom receive a money allowance from their parents and it is rare for young teenagers to earn pocket money, as jobs are assigned and controlled by the state or community. Thus, young people are not likely to have the money to buy cigarettes until they become wage earners.

The Chinese smoke fewer cigarettes compared to westerners. A limiting factor which has kept tobacco consumption low in China compared to other countries is the cost of cigarettes relative to income. Chinese domestic low grade



Source: Li et al, 1983

FIGURE 3—Per Cent of Shanghai Males Abstaining from Smoking by Duration of Longest Previously Reported Abstinence, Compared to Charleston Shipyard Workers

cigarettes cost thirty cents per pack, and better grades cost almost one yuan* per pack. The salary of Chinese wage earners averages about one hundred yuan per month in cities, and less in rural areas. Since about one-half of the salary goes for food, even one yuan adds up to a significant portion of the money available. The rising standard of living could lead to a rapid increase in cigarette consumption. Moreover, an increase in consumption could result in many social smokers becoming addicted.

Based on the limited data from Shanghai and a study of 848 male smokers in the Charleston naval shipyard conducted by Li and associates,[12] it appears that western smokers, possibly due to heavier consumption, have a harder time quitting smoking compared to Chinese smokers. The parallel curves in Figure 3, comparing the American smokers to the Chinese male smokers, suggest that in their attempts to quit, Chinese smokers are more likely than Americans surveyed to succeed in giving up cigarette smoking for at least one day (96 per cent vs 93 per cent), as well as for one year or longer (19 per cent vs 11 per cent), before relapse.

The main reason the Chinese smoke is social. Hosts offer cigarettes to guests as a sign of hospitality and sincere friendship. People light up as a gesture in accepting the hospitality and friendship. Government education to reduce smoking must focus on social needs of the people. It must promote role models for refusing cigarettes and create an environment in which smoking is seen as an anti-social behavior.

Paradoxically, because China is a planned society, the government is in a position to outlaw the production and sale of cigarettes, if there is political will to replace the revenue lost in banning tobacco production and manufacturing. Arguments for strategies for smoking control and long range planning for the diversion and substitution of tobacco crop have appeared in the *People's Daily*,[6] an official organ of the state and the most influential newspaper in China. Recently, Beijing took the major step of declaring 100 public places in the city smoke free, including hospitals, public transportation, and sports arenas.[13] Beijing's action in anti-smoking legislation could serve as a role model for the rest of the country.

* The exchange rate for Chinese currency (yuan) to the US dollar is approximately 1:3.7.

ACKNOWLEDGMENTS

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Advocacy Institute Issues Summary on Effects of Cigarette Smoking and Radon Exposure

The Washington-based Advocacy Institute, through its Smoking Control Advocacy Resource Center (SCARC), issued an "action-alert" news release in mid-October summarizing the effects of cigarette smoking on radon exposure and encouraging tobacco control advocates to seize the opportunity to "increase the public's perception of tobacco as an agent of disease that multiplies the hazards of radon and other environmental hazards." The message of the news release-summary is excerpted and paraphrased below:

On September 12, 1988, the US Environmental Protection Agency (EPA) issued a national health advisory concerning the effects of radon exposure. Radon, an invisible and odorless gas that is formed when uranium in soil and rocks decays, can produce high levels of radioactivity. According to EPA, the toxic action level occurs when the level of radon reaches 4 pci (picocuries) per liter of air—an amount equivalent to smoking half a pack of cigarettes a day.

EPA has estimated that eight million homes in the US may have potentially hazardous radon levels. Cigarette smoking, particularly in a radon-exposed house, would therefore pose a significant health threat to smokers and nonsmokers alike. In addition, EPA has stated:

- The cancer risk of smoking more than a pack of cigarettes a day is equivalent to spending 75% of the time in a house with a radon level that exceeds the health protection standards.
- Of the 20,000 deaths per year from radon-related lung cancer, 15,000 (75%) are smokers. Of the 140,000 total lung cancer deaths in America each year, smoking is responsible for 119,000.
- The risks for smokers exposed to radon is 15 times greater than for nonsmokers.
- Tobacco smoke acts as a carrier which draws radon particles from the air deep into the lungs. Therefore, smoking in a radon-exposed house increases the chances of radon-induced lung cancer for all persons in contact with the smoke.

According to William Mills, former Radiation Division Chief at EPA, "The message is, if you want to decrease your risk from radon, stop smoking; it's that simple."

The Advocacy Institute suggests that "Tobacco control advocates need to seize this occasion to convert the issue from a radon dilemma to a smoking problem." Some of the methods they suggest for accomplishing this include:

- Contact local radon testing programs to urge that they include surveying for smokers in the households they test. Ask that their inspectors stress "no smoking" as one easy and immediate corrective action.
- Contact local health reporters to emphasize tobacco as the agent of disease which multiplies the hazards of radon.

For further information about the activity, contact: Advocacy Institute, Smoking Control Advocacy Resource Center, 1730 Rhode Island Avenue, NW, Suite 600, Washington, DC 20036-3118. Tel: 202/659-8475.

Fractures and Lifestyle: Effect of Cigarette Smoking, Alcohol Intake, and Relative Weight on the Risk of Hip and Forearm Fractures in Middle-Aged Women

David Hemenway, PhD, Graham A. Colditz, MB, BS, Walter C. Willett, MD, Meir J. Stampfer, MD, and Frank E. Speizer, MD

Abstract: Cigarette smoking, alcohol consumption and low relative weight are often cited as risk factors for osteoporosis. In a prospective cohort study of 96,508 middle-aged nurses 35 to 59 years of age we found that smoking was not a risk factor for hip and forearm fracture. Women who drank more than 15 grams of alcohol per day and whose relative weight was less than 21 kg/m^2 were at increased risk of fractures, but these risk factors were not independent. Only the combination of alcohol intake and thinness substantially increased the likelihood of fracture. The low weight women consuming more than one drink per day comprised but 4 per cent of our population of middle-class women and sustained 6 per cent of the fractures. (*Am J Public Health* 1988; 78:1554–1558.)

Introduction

As with many other medical problems, osteoporosis has been associated with various lifestyle variables such as cigarette smoking,[1–4] high alcohol consumption,[4–6] and relative weight.[1,4,7] The relation between fractures and smoking and drinking has not been clearly documented,[8–12] but the association between fractures and relative weight is clear: a retrospective cohort study[13] and various case-control studies[8–10,14] have consistently shown that thin women are more likely to suffer hip and forearm fractures than their more obese counterparts.

We employed a prospective approach to investigate risk factors for fracture. With data from the Nurses' Health Study, a large prospective cohort study of middle-aged women with nursing training, we estimated the effect of cigarettes, alcohol, and obesity on the risk of fractures of the hip and forearm.

Methods

In 1976, 121,700 female registered nurses, ages 30 to 55, living in 11 large US states, completed a mailed questionnaire on known and suspected risk factors for cancer[15] and coronary heart disease.[16] Based on a subsample of the cohort, we estimate that at least 98 per cent are White. Every two years, a follow-up questionnaire has been sent to update and add to the information about these women and to record the occurrence of cancer, heart disease, and other major health problems.

In 1980, 103,000 women responded. From these we excluded women with a history of cancer (other than non-melanoma skin cancer) and the 4 per cent of women who did not answer either the 1982 or 1984 questionnaire leaving a total of 96,508 nurses. We used information on age, weight, height, smoking, and alcohol from the 1980 questionnaire. In 1982 and 1984, we asked participants whether they had sustained a fracture to the hip or forearm in the previous two years. Women responding yes on either questionnaire were considered cases if their fracture occurred after the return of the 1980 questionnaire and before June 1, 1984. To assess the validity of this self-report, we asked a random sample of 50 women who reported fractures on the 1982 questionnaire for permission to review the relevant medical records. Forty-three women responded and confirmed their prior self-report, and 33 gave us permission to review the records. All medical records obtained confirmed the self-report.[17]

Cumulative incidence rates were determined for fractures occurring during the four years of follow-up (1980–84). Relative risks (cumulative rate ratios) were calculated as a measure of association. Using the self-reported age on the 1980 questionnaire, we divided the women into five age categories of roughly equal size: ages 35–39 (22 per cent of the women); 40–44 (20 per cent); 45–49 (21 per cent); 50–54 (20 per cent); 55–60 (17 per cent). Respondents were also divided into five smoking categories: never smoked (44 per cent); ex-smoker (27 per cent); one to 14 cigarettes per day (8 per cent); 15 to 24 cigarettes (12 per cent); more than 25 (8 per cent). Less than 2 per cent of the women failed to respond to this question.

We estimated the number of grams of alcohol intake per day. Respondents were asked about their average use in the past year of beer (bottles or cans), wine (glasses), and liquor, i.e., whiskey, gin, etc. (drinks). They were instructed to select one of nine categories for each type of alcohol: 6+ per day; 4-6 per day; 2–3 per day; 1 per day; 5–6 per week; 2–4 per week; 1 per week; 1–3 per month; almost never. The alcohol content was estimated to be 13.2 grams for a bottle or can of beer, 10.8 grams for a glass of wine, and 15.1 grams for a drink of liquor. Total alcohol intake for each individual was computed as the sum of the contribution from beer, wine, and liquor. We divided the women into five groups: total alcohol intake per day = 0 (30 per cent); between 0 and 1.4 grams/day (12 per cent); from 1.5 to 4.9 grams/day (20 per cent); from 5 to 15 grams (19 per cent); and greater than 15 grams (13 per cent). (Five grams per day is the equivalent of about three drinks per week). Seven per cent of the women did not provide plausible responses to the diet component of the 1980 questionnaire, and thus were not categorized for alcohol intake. Alcohol intake responses have been validated as part of the overall validation of the Nurses' Health Study food frequency questionnaire.[18,22] Questionnaire data were highly correlated with alcohol intake recorded during four one-week diet periods ($r = 0.90$), and were reproducible over one year ($r = 0.90$).

Relative weight was calculated as weight in kilograms

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TABLE 1—Hip and Forearm Fractures 1980–84 (4 years) by Age

	Age (years)				
Age in 1980	30–35	40–44	45–49	50–54	55–60
No. of Cases	94	84	155	301	291
No. of Women	21,334	19,501	20,118	19,429	16,126
Incidence (cases/10,000 person-years)	11.0	10.8	19.3	38.7	45.1

divided by height in meters squared (Quetelet's index) and categorized as: less than 21 kg/m^2 (21 per cent of the women); 21 to 23 (24 per cent); 23 to 25 (20 per cent); 25 to 29 (20 per cent) over 29 (14 per cent). Six per cent of the women did not provide sufficient information and were not categorized. A relative weight of under 21 kg/m^2 corresponds to a 5'4" woman weighing less than 122 pounds. A relative weight of more than 29 kg/m^2 corresponds to a 5'4" woman weighing more than 169 pounds.

Weight data for the nurses also have been collected as part of a validation study which showed that self-reported weight was highly correlated with measured weight (r = 0.96).[15]

Results

Slightly under 1 per cent of the respondents (925 of 96,508) reported a fracture to the hip or forearm in the period from the date of return of the 1980 questionnaire to June 1, 1984.

Like other observers, we found a strong increase in fracture rates with increasing age (Table 1). Women in the 55–60 age group had over four times the likelihood of experiencing a fracture in the succeeding four years compared to women ages 35–44.

We observed no relation between smoking categories and likelihood of fracture either overall or within any five-year age category (Table 2). There was some suggestion of an association between alcohol consumption and fractures (Table 3). Among women consuming 15 or more grams of alcohol per day, compared to abstainers and light drinkers, the relative risk was 1.24 (95 per cent confidence interval 1.04, 1.48) (Table 3).

Compared to those of higher relative weights, the thinnest women (<21 kg/m^2) were at increased risk for fracture. (Table 4). For women with relative weights above 21 kg/m^2, the age-adjusted risk of fractures did not vary by weight. The increased risk observed for lean women was limited to women over 50 years of age.

We analyzed the interrelation between thinness and alcohol intake characteristics (Table 5). Women who drank more than 15 grams of alcohol per day were not at increased risk unless they were thin, and women who were thin were not at increased risk unless they drank more than 15 grams of alcohol per day. The low weight women who drank most heavily had an age-adjusted risk of fracture 70 per cent above the rest of the women in the study (relative risk = 1.73, 95 per cent CI = 1.30, 2.29). In terms of absolute risk, the low weight-heavier drinkers accounted for under 4 per cent of the study population and reported 6 per cent of the fractures. We observed no interaction between relative weight, cigarette smoking, and risk of fractures in this population.

Discussion

A recent consensus conference argued for additional "observational and epidemiologic studies to determine the impact of multiple and behavioral factors on . . . fracture frequency. Such studies could be conducted by appropriate additions to existing population-based studies."[19] In accord with that recommendation, we examined the relationship between the risk of hip and forearm fracture and three lifestyle characteristics: cigarette smoking, alcohol consumption, and relative weight. The study was large—prospective data were collected on close to 100,000 nurses—so any substantial relationship would be detectable. However, the fact that the women in the study were middle-aged (no one was over 65 in 1980) may limit the generalizability of the results.

While the data are based on recall, the presumption that these trained nurses are reliable reporters has been corroborated by various validity checks.[15,17,18,20] In addition, the questionnaire data have been used to corroborate associations reported from other epidemiologic studies of risk factors for heart disease and cancer.[15,16,21–23]

We observed a marked increase in fracture rates with age. The absolute rates reported are probably slight under-

TABLE 2—Incidence of Fractures, by Age, Smoking (cases/10,000 person-years)

	Age (years)					Age-Adjusted Relative Risk (Never Smoked = 1.00)	95% Confidence Interval
Smoking Day	35–39	40–44	45–49	50–54	55–60		
Never Smoked							
Cases: 413							
Pop: 42,185	10.5	10.7	19.4	43.4	45.4	1.00	
Ex-Smoker							
Cases: 252							
Pop: 26,134	10.3	12.5	19.6	37.2	47.0	0.98	(0.84, 1.14)
1–14 Cigarettes							
Cases: 65							
Pop: 7,307	10.8	10.0	15.7	31.4	54.5	0.95	(0.71, 1.20)
15–24 Cigarettes							
Cases: 108							
Pop: 11,350	10.4	11.0	19.0	40.5	44.9	0.97	(0.79, 1.20)
25+ Cigarettes							
Cases: 76							
Pop: 7,876	17.7	7.5	19.7	36.9	47.0	0.99	(0.78, 1.25)

TABLE 3—Fracture Incidence by Age, Alcohol Intake (cases/10,000 person-years)

Alcohol Grams/Day	Age (years) 35–39	40–44	45–49	50–54	55–60	Age-Adjusted Relative Risk	95% Confidence Interval
None							
Cases: 277	10.5	11.3	22.0	37.9	42.8	1.00	
Pop: 28,668							
0 to 1.4 gms							
Cases: 98	8.6	12.4	17.9	37.5	40.5	0.95	(0.74, 1.16)
Pop: 11,441							
1.5–4.9 gms							
Cases: 178	10.5	10.7	24.1	40.1	46.8	1.06	(0.88, 1.27)
Pop: 18,588							
5.0–14.9 gms							
Cases: 166	11.4	8.7	17.5	36.8	49.4	0.98	(0.81, 1.19)
Pop: 17,912							
15 or more gms							
Cases: 151	16.2	10.8	11.0	51.6	56.9	1.16	(0.95, 1.41)
Pop: 12,792							
Under 15 gms							
Cases: 719	10.4	10.7	20.8	38.1	44.9	1.00	
Pop: 76,609							
15 or more gms							
Cases: 151	16.2	10.8	11.0	51.6	56.9	1.24	(1.04, 1.48)
Pop: 12,792							

estimates of the real rates. A few respondents may have had more than one fracture in the preceding two years, but only one fracture could be reported.

Our finding that cigarette smoking was not a risk factor for fractures agrees with the findings reported from case-control studies by Paganini-Hill, *et al*,[9] and Kreiger, *et al*,[10] for hip fracture, and by Williams, *et al*,[11] for forearm fracture. On the other hand, Alderman, *et al*,[8] and Williams, *et al*,[11] showed an association between smoking and hip fracture. Williams, *et al*,[11] found a particularly strong effect among thin women, an effect we did not find.

It is possible that smoking increases the risk of osteoporosis,[1–4] but did not alter the risk of fracture among these nurses because they were too young. As the population ages, and sustains more osteoporosis and fracture, we will be able to examine the question further.

Alcohol intake is a risk factor for most injuries, including falls.[24,25] Most fractures are due to falls.[26] Indeed, more than 90 per cent of the Nurses' Health Study population who described the circumstances of their fracture on the 1986 questionnaire (over 600 respondents) had fallen. Alcohol consumption may also have the long-term effect of reducing bone density.[4–6] In addition, it can have the immediate impact of evoking risk-taking behavior and decreasing balance, resulting in more falls and fractures. The diet assessment questionnaire used in 1980 inquired about average daily consumption over the preceding year, and we do not have the blood alcohol content at time of injury. We might therefore expect to observe less of an alcohol-injury relationship than case-control studies which measure blood alcohol content at time of medical treatment.[24,25]

We observed no differences in the risk of fracture between teetotalers and those consuming up to 15 grams of alcohol per day. Alcoholism has been cited as a substantial risk factor for hip and forearm fracture.[14] We had no measure of alcoholism, but those drinking the most in our population did have a 24 per cent higher likelihood of fracture. Since the heavy drinkers comprised less than 15 per cent of our study

TABLE 4—Fracture Incidence by Age, Relative Weight (cases/10,000 person-years)

Relative Weight Kg.m^2	Age (years) 35–39	40–45	45–49	50–54	55–60	Age-Adjusted Relative Risk (<21 relative weight = 1.00)	95% Confidence Interval
<21							
Cases: 192	11.8	10.6	16.6	64.7	49.5	1.00	
Pop: 19,363							
21–23							
Cases: 200	6.8	9.0	20.5	38.2	54.0	0.82	(0.67, 0.99)
Pop: 22,068							
23–25							
Cases: 186	17.3	11.1	19.9	38.6	41.0	0.85	(0.70, 1.04)
Pop: 18,297							
25–29							
Cases: 188	13.7	12.3	16.4	39.3	46.4	0.77	(0.63, 0.95)
Pop: 18,145							
29+							
Cases: 127	13.5	15.0	23.7	24.5	44.9	0.80	(0.64, 0.99)
Pop: 12,873							

TABLE 5—Fracture Incidence by Age, Alcohol Intake, Relative Weight (cases/10,000 person-years)

Relative Weight/Alcohol Intake	Age (years)					Age-Adjusted Relative Risk (Not Thin–Not Heavy Drinker = 1.00)	95% Confidence Interval
	35–39	40–44	45–49	50–54	55–60		
Not Thin*; Not Heavy Drinker**							
Cases: 582							
Pop: 60,369	10.5	11.1	21.5	35.4	45.4	1.00	
Not Thin; Heavy Drinker							
Cases: 100							
Pop: 9,451	11.6	11.2	11.3	39.3	53.1	1.01	(0.82, 1.25)
Thin; Not Heavy Drinker							
Cases: 136							
Pop: 15,752	10.1	8.9	18.4	55.2	44.8	1.09	(0.90, 1.31)
Thin; Heavy Drinker							
Cases: 51							
Pop: 3,270	24.8	9.7	10.3	99.8	74.9	1.73	(1.30, 2.29)

*Thin: relative weight less than 21 Kg/m^2
**Heavy Drinker: 15 or more grams/day of alcohol

population, even if their risk of fracture had been reduced to the level of non-drinkers, total reported fractures would have decreased only 3 per cent.

Relative weight may be associated with fractures for a variety of reasons. Heavier women, for example, may be larger boned, and have more bone mass to lose before the process of osteoporosis leads to critically low levels. Additional adipose tissue may provide more opportunity for estrogen production, and the fat itself may offer cushioning protection during a fall. Other studies of hip fracture,[9,14] and of forearm fracture,[8,14] as well as this study, find the thinnest women at somewhat higher risk for fracture.

Our finding that the combination of thinness and relatively heavy alcohol consumption dramatically increased the fracture risk (particularly among women over 50 years old) is provocative and deserves further research.

While this study has the advantage of being large and prospective, it suffers from several limitations. It included few elderly women, who are at greatest risk for fracture and subsequent mortality. Second, the 1982 and 1984 follow-up questionnaires combined hip and forearm fractures together, although they may be of different etiologies.[19] Finally, in this analysis we examine only four independent variables: age, cigarette smoking, alcohol consumption, and relative weight, although there are many other potential confounders.

Nevertheless, a conclusion that does emerge from the analysis is how little is explained by the three lifestyle variables examined. Most similar-aged women in the study—whether thin or fat, smoking or non-smoking, teetotalers or drinkers—generally had an equivalent chance of sustaining a hip or forearm fracture. Only the thin drinkers had a substantially higher risk, and they represented only 4 per cent of the population and 6 per cent of the fractures. The evidence suggests that otherwise beneficial habits such as avoidance of smoking and obesity may not significantly reduce the risk of fracture among middle-aged women.

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Ford Foundation and Harvard University Honor Government Ingenuity

Ten creative state and local government projects addressing some of today's most difficult and seemingly intractable public policy problems were recently named winners in the 1988 Innovations in State and Local Government Awards Program, sponsored by the Ford Foundation and the John F. Kennedy School of Government at Harvard University.

Ranging from a welfare-to-work program to a housing integration initiative in Cleveland, Ohio's suburbs, to a project to house the homeless in San Diego, California, the winners received $100,000 grants from the Ford Foundation at a ceremony held September 30 at the National Press Club in Washington, DC. They include five statewide and four municipal programs, as well as one county-sponsored project selected from among 970 entries. The field was narrowed to 25 finalists last June. The 1988 winners are:

CALIFORNIA—*The City of San Diego's Single-Room Occupancy Residential Hotel Program* reverses the decline of privately-owned housing for the city's low-income population, and eases the crisis of homelessness by encouraging the private sector rehabilitation and construction of single-room occupancy hotels (SRO's).
COLORADO—*The City of Fort Collins' Land Development Guidance System (LDGS)* creates a new, flexible guide for urban planning and development by incorporating issues of land use, design, density, and location into a single decision-making process.
ILLINOIS—*Project Match of the Department of Public Aid* encourages welfare recipients to enter the workforce by helping them return to school, obtain job training, find and keep jobs, and advance in a career.
KENTUCKY—*Court of Justice of the Commonwealth of Kentucky* has replaced court reporters and audiotapes with videotapes in trial courts across the state, providing instantaneous records of court proceedings for lawyers and juries, and lowering the cost of justice.
KENTUCKY—*The Kentucky Department of Education's Parenthood and Child Education Program* stops the intergenerational cycle of illiteracy in rural areas by combining adult education with early childhood development.
MASSACHUSETTS—The state's *Industry Action Program* promotes competitiveness and job preservation in mature industries by providing the resources necessary to address problems in a specific region.
MISSISSIPPI—*The City of Tupelo's Municipal School District* is working in partnership with private citizens, business, city government, and local foundations to make Tupelo one of the top 10 public school systems in the nation.
MISSOURI—*The St. Louis County Police Department* has reduced the amount of time officers spend filing reports by some 80 per cent, with the first comprehensive automated system encompassing reporting, management, and investigation.
OHIO—*The Racial Integration Incentives*, developed by the City of Shaker Heights Department of Community Services and financed by The Fund for the Future of Shaker Heights, along with two other organizations, promotes racial integration in a cluster of area suburbs in eastern Cuyahoga County by attracting home buyers to areas where their race is underrepresented.
VERMONT—*The State of Vermont Department of Libraries* has created a Statewide Library Automation Project to make the state's public and private library collections available to every citizen.

The $4.5 million Awards program, a three-year initiative now in its third and final year, has attracted 3,000 entries from across the United States since 1986. The Ford Foundation plans to continue the program in a new expanded form beginning in 1990, in cooperation with selected universities in several regions throughout the country. For further information, contact the Ford Foundation, 320 East 43 Street, New York, NY 10017. Tel: (212) 573-5000.

Drug Abuse Discharges from Non-Federal Short-Stay Hospitals

JOSEPH C. GFROERER, BA, EDGAR H. ADAMS, ScD, AND MARY MOIEN, MS

Abstract. An analysis of inpatient drug abuse cases was done using the National Hospital Discharge Survey (NHDS). An estimated two million discharges with a drug abuse diagnosis occurred in non-federal short-stay hospitals during 1979–85, a figure which is believed to be an underestimate. Compared to other hospital inpatients, drug abuse inpatients are more likely to be male, ages 15–44, and other than White race. Increases in hospital use for drug abuse treatment were found to have occurred between 1979 and 1985, with discharge rates per 10,000 population increasing from 3.1 to 6.0 for drug dependence and from 3.8 to 7.7 for nondependent drug abuse. Concurrent increases in availability of hospital-based inpatient drug and alcohol treatment programs and insurance coverage for drug abuse treatment were found to have occurred during the same period. (*Am J Public Health* 1988; 78:1559–1562.)

Introduction

Data on people receiving treatment for drug abuse have been collected and analyzed for a number of years by the National Institute on Drug Abuse (NIDA).[1–6] Nevertheless, there are no published national data on patients treated for drug abuse as inpatients in short-stay hospitals.

The purpose of this paper is to analyze drug abuse treatment data from the National Hospital Discharge Survey (NHDS) for the period 1979–85.

In addition, this study will evaluate the usefulness of the NHDS as a data source for drug abuse descriptive epidemiology and surveillance. Because drug abuse is an illicit deviant behavior, data on trends and changes in drug use patterns are difficult to obtain. The availability of data from a variety of sources can serve to provide corroboration of findings and occasionally may highlight regional variability. The NHDS population includes drug abusers referred to hospitals from drug treatment programs and private physicians, as well as people coming into hospital emergency rooms who are admitted as inpatients. NIDA's Drug Abuse Warning Network (DAWN) indicates that over 40 per cent of emergency room episodes involving drug abuse result in admission to the hospital.[5] Because of this overlap, comparisons between DAWN and NHDS will be discussed.

Methods

Data Source

The NHDS is conducted by the National Center for Health Statistics (NCHS) on a continuous basis. The data are obtained from the face sheets of a sample of inpatient medical records that are abstracted from a national sample of non-federal short-stay general and specialty hospitals in the United States. The sample is updated periodically to account for changes in the universe, such as the opening of new hospitals. Approximately 195,000 medical records from 407 participating hospitals were sampled in the 1985 NHDS, representing approximately 0.5 per cent of all discharges from non-federal short-stay hospitals in the US in 1985.[7,8] The response rate for hospitals has always been over 80 per cent in the NHDS. Sample data are weighted to provide nationally representative estimates of hospital discharges. Estimates of relative standard errors are provided to users of the data in charts shown in NHDS reports which approximate relative sampling errors for a wide variety of estimates. Over a seven-year period, the approximate relative standard error of an estimate of 100,000 discharges would be 9 per cent, and for an estimate of 20,000 discharges it would be 13 per cent.

Information on patient characteristics, admission and discharge dates, expected method of payment, geographic region of hospital, medical diagnoses and procedures are collected. Diagnosis and surgical and nonsurgical procedures are coded using the International Classification of Diseases (ICD) system. Since 1979, *The International Classification of Diseases—9th Revision—Clinical Modification* (ICD-9-CM)[8] has been used. Up to seven ICD-9-CM diagnostic codes are produced from each abstract record, with the first-listed code usually representing the principal diagnosis.

Definitions and Analytic Methods

Drug abuse discharges were defined as those with any listing of the following diagnostic categories:

ICD-9-CM Code	*Description*
304	Drug dependence
305.2–305.9	Non-dependent abuse of drugs, excluding alcohol abuse and tobacco use disorder
965.0, 965.8, 967, 968.5, 969	Poisoning by drugs, including only opiates and related narcotics, pentazocine, sedatives and hypnotics, cocaine, and psychotropic agents

The "poisoning by drugs" codes were included in the drug abuse definition because these codes involve many drugs which are known to be widely abused such as opiates, pentazocine, sedatives and hypnotics, cocaine, antidepressants and tranquilizers. Drug abuse discharges that included an ICD-9-CM code for suicide (codes E950–E958) were included in this study. Suicide attempts are known from DAWN to be involved in many emergency room episodes for sedatives and tranquilizers. Specifically excluded were poisoning by aspirin, acetaminophen, anticonvulsants, and anesthetics (except cocaine).

Also excluded from the drug abuse definition were drug psychoses cases (ICD-9-CM code 292) that did not specifically mention one of the above listed drug categories. Analysis of these cases showed that most involved a "correct drug properly administered" (ICD-9-CM codes E930–E949) as causing the psychoses.

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TABLE 1—Drug Abuse Discharges from Hospitals by Diagnostic Category, 1979–85

Diagnostic Category	ICD-9-CM Codes	Discharges in Thousands	
		First Listed	All Listed
Total drug abuse		829	1,998
Drug dependence	304	294	718
Opiates	304.0	74	177
Cocaine	304.2	14	41
Marijuana + hashish	304.3	25	99
Unspecified drug dependence	304.9	140	313
Non-dependent drug abuse (excluding alcohol, tobacco)	305.2–305.9	213	903
Marijuana	305.2	7*	51
Opiates	305.5	12	45
Cocaine	305.6	12	57
Unspecified drug abuse	305.9	173	742
Poisoning by drugs	965.0, 965.8, 967, 968.5, 969	322	449
Opiates and related narcotics	965.0	18	36
Pentazocine	965.8	12	19
Sedatives and Hypnotics	967	74	107
Cocaine	968.5	**	9*
Psychotropic agents	969	215	302

*Based on small sample (30–59 cases).
**Sample size too small to be reliable (less than 30 cases).

Data analysis involved tabulation of drug abuse discharges from 1979 to 1985. Weighted data were used for all tables, so that the data represent estimates of discharges from all non-federal short-stay hospitals in the United States. Only differences that are statistically significant at the .05 level are specifically mentioned in the text.

Results

An estimated 2 million hospital discharges involving drug abuse occurred during 1979–85, with 829,000 having drug abuse as the principal diagnosis. Table 1 shows the distribution of these discharges by specific diagnosis. A large proportion of drug abuse discharges on the NHDS file have the drug unspecified: 48 per cent of first-listed drug dependence diagnoses, and 81 per cent of first-listed non-dependent drug abuse diagnoses. Thus, comparison of population rates and totals among specific drug types is not possible.

Table 1 also shows that when drug poisoning is listed as a diagnosis, it usually is listed first (72 per cent of the 449,000 discharges), while this is not the case for the other drug abuse categories (41 per cent for drug dependence and 24 per cent for nondependent drug abuse). However, analysis of these other drug abuse categories suggests that even when drug abuse was not listed first, it was usually directly related to the first-listed diagnosis. In these cases, alcohol abuse or psychiatric disorders were usually listed first.

Drug abuse accounted for 0.8 per cent of an estimated 263 million discharges (excluding newborn infants) during 1979–85. Table 2 shows that drug abuse was more likely to be involved in discharges for males (54 per cent vs 40 per cent), for people ages 15–44 (76 per cent vs 40 per cent), and for races other than White (22 per cent vs 14 per cent).

While drug dependence and non-dependent drug abuse discharges appear to be very similar in terms of demographics, drug poisoning discharges are more likely to be female, more likely to be under age 15 or over age 44, and more likely to be White.

The drug poisoning discharges also are more likely to involve suicide attempts. While 18 per cent of drug poisoning discharges mention suicide (ICD codes E950–E958), less than 1 per cent of drug dependence and non-dependent drug abuse discharges do so. While keeping in mind the underreporting involved because most NHDS discharges for drug abuse are coded as drug unknown, the demographic characteristics of patients discharged for opiates, cocaine, and marijuana show interesting differences. Only 13 per cent of the opiate discharges were under age 25, while 32 per cent of cocaine and 58 per cent of marijuana discharges were under age 25. The marijuana discharges were more likely to be White than were opiate or cocaine discharges (73 per cent vs 47 and 50 per cent). Characteristics of patients discharged for poisoning by barbiturates, benzodiazepine tranquilizers, and antidepressants showed patterns similar to those for the combined drug poisoning category.

Drug abuse discharge rates are highest for 25–34 year-

TABLE 2—Demographic Distribution of All-listed Drug Abuse Discharges and Non-Drug Abuse Discharges, 1979–85

Characteristics	No Drug Abuse Diagnosis	Drug Abuse Diagnosis	Drug Dependence (304)	Non-dependent Drug Abuse (305.2–305.9)	Poisoning by Drugs (965.0 965.8, 967 968.5, 969)
Number of discharges (in thousands)	260,718	1,998	718	903	449
Sex					
% Male	40	54	61	59	38
% Female	60	46	39	41	62
Age (years)					
% Under 15	9	3	2	2	7
% 15–24	14	26	25	29	25
% 25–34	16	35	37	38	27
% 35–44	10	15	16	15	15
% 45+	50	21	21	16	26
Race					
% White	77	67	63	66	78
% Other	14	22	24	24	13
% Unknown	9	11	13	10	9

NOTE: Excludes newborn infants.

TABLE 3—Trends in All-listed Drug Abuse Discharge Rates by Major Diagnostic Category, Sex, and Age (rates per 10,000 population)

Major Diagnostic Category	1979	1981	1983	1985
Drug Dependence (ICD Code 304)				
Total	3.1	4.3	4.6	6.0
Sex				
Male	3.9	5.1	5.7	8.0
Female	2.4	3.5	3.5	4.2
Age (years)				
Under 15	**	**	**	**
15–24	4.5	6.4	6.0	7.4
25–34	6.7	9.3	10.2	14.4
35–44	3.4*	6.6	5.7	8.4
45+	2.5	2.5	3.1	3.7
Non-dependent Abuse of Drugs (ICD Codes 305.2-305.9)				
Total	3.8	4.6	6.3	7.7
Sex				
Male	4.3	5.4	7.6	9.8
Female	3.2	3.8	5.0	5.7
Age (years)				
Under 15	**	**	**	**
15–24	6.6	8.2	10.1	13.4
25–34	7.3	9.4	14.5	18.4
35–44	3.4	4.7	8.0	9.6
45+	2.8	2.6	3.0	2.9
Poisoning by Drugs (ICD Codes 965.0, 965.8, 967, 968.5, 969)				
Total	3.0	2.8	2.4	3.3
Sex				
Male	2.1	2.1	1.6	2.7
Female	3.7	3.4	3.1	3.8
Age (years)				
Under 15	1.2*	**	**	1.1*
15–24	4.6	4.2	3.3	4.5
25–34	4.2	4.6	3.6	5.5
35–44	2.4*	4.0	2.8*	3.4
45+	2.8	1.9	2.4	2.8

*Based on small sample (30–59 cases).
**Sample size too small to be reliable (less than 30 cases).

olds in all three diagnostic groups (Table 3). Males have higher rates for drug dependence and nondependent abuse while rates for drug poisoning are higher for females. Discharge rate trends by diagnostic group show that, while rates roughly doubled between 1979 and 1985 for both drug dependence and non-dependent drug abuse, rates for poisoning by drugs declined somewhat from 1979 to 1983, then increased in 1985 (Table 3). These trends occurred for both males and females, and for most age groups. Trends for specific drug types suggest very large increases in discharges for opiates, cocaine, and marijuana and relatively stable rates for poisoning by barbiturates, benzodiazepine tranquilizers, and antidepressants.

Discussion

The usefulness of NHDS data on drug abuse treatment is limited by both sampling and non-sampling errors. Because of the relatively low rate of discharge for drug abuse (compared to heart disease, cancer, or delivery, for example), sampling errors are unacceptably high for some subgroup estimates. To perform analysis of small subgroups, it is therefore necessary to combine data years; thus, analysis of annual data, such as for trends, can only be done for large population subgroups or for combined diagnostic categories. These approaches were employed in the tabulations shown in this paper.

The use of NHDS published relative standard error (RSE) charts also has its limitations. While these charts represent an approximation of the RSE for a wide variety of estimates, it is possible that errors for estimates of drug abuse discharges are somewhat larger than errors for other diagnoses because of the distribution of the drug abuse discharges among the hospitals in the sample. If this is the case, the RSE charts would underestimate the actual sampling errors associated with drug abuse discharge estimates.

The major non-sampling error problem with the NHDS drug abuse data involves the diagnostic coding. For a variety of reasons, it is likely that underreporting of drug abuse diagnoses occurs. First, the stigma associated with a patient having a drug abuse diagnosis may result in physicians in some cases not including drug abuse information. This practice could also be related to insurance coverage in that patients with no coverage for drug abuse treatment may be given other diagnoses that are covered by insurance. Also, physicians may not be sufficiently trained to recognize drug abuse, especially in comorbid conditions. In one recent study, physicians identified only 40 per cent of patients who suffered from alcohol or other substance abuse.[9]

Finally, since the NHDS abstractors primarily use only the face sheet, some diagnoses may not be identified, even if they are in the body of the medical record. Thus, it is likely that the proportion of hospital discharges related to drug abuse was larger than the 0.8 per cent estimated from the NHDS for 1979–85.

The use of face sheets may also lead to inaccurate coding of principal diagnoses, as was shown in an NHDS reliability study.[10] In that study, for patients found to have mental disorders, of which drug dependence and nondependent abuse are subgroups, 81 per cent of principal diagnoses were identified by NHDS, compared to 90 per cent of all-listed mental disorders that were identified by NHDS.

Thus the all-listed data are a better indicator of drug abuse involvement in the hospitalized population. A separate analysis, not tabulated here, showed that very little double-counting of discharges occurs, since less than 5 per cent of the discharges had diagnoses in more than one of the three major drug abuse diagnostic categories. However, specific drug type categories do involve some double-counting, especially the cocaine category. About half of the estimated 98,000 discharges for cocaine dependence or non-dependent abuse also had diagnoses of marijuana or opiate abuse, a finding which is consistent with other studies of cocaine treatment patients.[2,6]

It must also be stressed that NHDS estimates reflect counts of discharges, not people. It is unknown to what extent people have multiple inpatient hospital episodes during a year or a seven-year period.

Despite its limitations, the NHDS appears to be a useful source of drug abuse treatment data. As an ongoing nationally representative survey, it can be used to monitor trends in the use of inpatient hospitalization for drug abuse treatment. Patient characteristic and diagnostic information available from NHDS provide descriptive epidemiologic statistics for a population of drug abusers receiving the most intense and expensive type of care.

A comparison of NHDS and DAWN data[6] indicated generally consistent patterns in terms of patient characteristics (age, sex, and race) and trends. There are several major

differences in coverage between the two surveys, however. DAWN collects data from a non-random sample of hospital emergency rooms primarily in 27 major metropolitan areas, with a supplemental national panel of emergency rooms outside those areas. Although over 40 per cent of the drug-related emergencies are admitted as inpatients and thus become part of the NHDS universe, over half are released without being admitted.[5] Similarly, many NHDS drug-related cases are not admitted through the hospital emergency room and are therefore not eligible for DAWN. Hospital admissions may reflect chronic conditions rather than the acute conditions generally seen in emergency rooms. Another difference in the two surveys is that the NHDS includes some short-stay specialty hospitals, while DAWN includes only general hospitals.

Nevertheless, similarities in the DAWN and NHDS data suggest that despite the large proportion of NHDS cases with unknown drug, the discharges that do specify drug provide useful information on the characteristics of drug abuse discharges. However, for barbiturates, benzodiazepine tranquilizers, and antidepressants, DAWN metropolitan area and national panel data indicate decreases while NHDS data do not do so. Since NIDA is currently implementing a nationally representative sample in DAWN, comparisons between NHDS and DAWN will be much more informative in the near future.

An interesting finding from the NHDS data is the sharp increase in utilization of hospitals for drug abuse treatment between 1979 and 1985, a period during which the prevalence of drug use in the US population was stable or declining, with the exception of cocaine use.[11] This increase in hospital discharges may be related to increasing availability of hospital care and the increasing availability of insurance coverage for drug abuse treatment. Availability of hospital inpatient treatment for drug abuse can be measured by the number of community hospitals which have alcohol/chemical dependency inpatient units and by the number of beds in those units. The trend in availability of these units[12–15] closely parallels trends in discharge rates for drug dependence and non-dependent drug abuse (Table 3). A Department of Labor survey of medium and large companies[16,17] estimated that 36 per cent of workers in these companies had drug abuse treatment coverage in 1982, while 59 per cent had coverage in 1985. Also, increase in the demand for treatment may be occurring, as the drug-using cohorts of the 1960s and 1970s become older and the long-term health consequences of chronic drug use begin to emerge. Moreover, data from clients in drug abuse treatment facilities suggest a lag period of several years between first use of drugs and first entry into treatment.[18,19] Thus, an increase in treatment demand would be expected to follow by several years the increases in prevalence of drug use which occurred in the 1970s. Trends toward more dangerous patterns of use (such as smoking cocaine and polydrug use) in recent years may also be influencing treatment demand. Finally, it is also possible that physicians are becoming more adept at recognizing patients' drug abuse problems.

In summary, the NHDS drug abuse data are useful as a descriptive epidemiologic surveillance tool which when used with other data can serve to increase our understanding of drug abuse problems and shifting treatment demand. Users should be aware of the limitations of the data, however, particularly the likelihood that drug abuse diagnoses are underreported, with the implication being that the NHDS data underestimate the actual level of hospital discharges involving drug abuse.

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Prevention of Lower Extremity Stress Fractures: A Controlled Trial of a Shock Absorbent Insole

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Abstract: A prospective controlled trial was carried out to determine the usefulness of a viscoelastic polymer insole in prevention of stress fractures and stress reactions of the lower extremities. The subjects were 3,025 US Marine recruits who were followed for 12 weeks of training at Parris Island, South Carolina. Polymer and standard mesh insoles were systematically distributed in boots that were issued to members of odd and even numbered platoons. The most important finding was that an elastic polymer insole with good shock absorbency properties did not prevent stress reactions of bone during a 12-week period of vigorous physical training. To control for the confounding effects of running in running shoes, which occurred for about one and one-half hours per week for the first five weeks, we also examined the association of age of shoes and cost of shoes with injury incidence. A slight trend of increasing stress injuries by increasing age of shoes was observed. However, this trend did not account for the similarity of rates in the two insole groups. In addition, we observed a strong trend of decreasing stress injury rate by history of increasing physical activity, as well as a higher stress injury rate in White compared to Black recruits. The results of the trial were not altered after controlling for these factors. This prospective study confirms previous clinical reports of the association of stress fractures with physical activity history. The clinical application of a shock absorbing insole as a preventive for lower extremity stress reactions is not supported in these uniformly trained recruits. The findings are relevant to civilian populations. (*Am J Public Health* 1988; 78:1563–1567.)

Introduction

Stress fractures and stress reactions of the lower extremities have long been considered a hazard of military life and are frequently associated with road marching or high intensity physical training. The occurrence of these injuries has been noted for many decades,[1–5] but only since the mid-1970s have studies with rate calculations been published, yielding a better idea of the serious nature of this problem. Eight studies since 1974 have documented that the stress fracture rate in military recruits varies from 1 to 4 per cent in males and from 10 to 15 per cent in females.[6–13] Injuries of this type, particularly if they occur in the femur, have a protracted hospital course, can be permanently disabling, and lead to a higher rate of medical discharge in affected trainees.

The ideal military training program would allow trainees to pass through a period of bone resorption activity, followed by bone replacement activity, without suffering a stress fracture or periosteal stress reaction.[13] Prevention in military training has therefore focused on reducing the intensity of training in the first few weeks, the period of maximal bone resorption, particularly for trainees who are not physically fit at entrance. Current US Army training policy requires unfit trainees to be exercised more gradually during the first three weeks.*

One characteristic of the training environment that can be modified easily is the amount of cushioning provided by footwear during training. Lack of cushioning in shoes has been hypothesized for some time as a cause of running injuries.[14–17] Consequently, running shoes are permitted to be worn during training runs in the military in place of the military combat boot.[18] The military combat boot was not designed to be shock absorbent, but to protect the foot and ankle from inversion injuries. An evaluation of Army combat boots found them to be inferior to running shoes in impact protection, flexibility, and energy retrieval.[19] In spite of these shortcomings, many hours of running, marching, calisthenics, and drill take place in combat boots. The simplest means for enhancing impact protection of the boot is to place within it a shock-absorbing insole. If it were to prove efficacious in preventing lower extremity injuries, such an insole could easily be added to the military footwear currently issued to each recruit.

The Sorbothane insole (IEM Orthopaedics, Aurora, Ohio), a viscoelastic polymer, was manufactured to absorb up to 95 per cent of the impact stress transmitted from the ground to the lower extremities. The following report describes a 1985 trial of the effect of Sorbothane insoles, as compared to a standard insole made of mesh, in prevention of stress fractures and stress reactions of bone.

Methods

Subjects

The subjects of this study were 3,025 United States Marine basic training recruits, with a mean age of 20.0 years ± .02, range 18 to 41, who underwent 12 weeks of basic training at the Marine Training Center at Parris Island, South Carolina. Sixty-seven per cent of the trainee population was White non-Hispanic, 24.6 per cent was Black non-Hispanic, 5.6 per cent was Hispanic, 0.7 per cent was Native American, and 1.5 per cent was of other ethnic origin.

Questionnaire

Recruits completed a questionnaire after hearing each item read aloud by an instructor to ensure comprehension. Questions were included which determined the recruits' usual physical activity status (very active, active, average, not very active, inactive) prior to their arrival at Parris Island, and the age, condition, and cost of their running shoes used during training.

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*US Army Training and Doctrine Command: Physical Training Program of Instruction (letter). October 1985, Fort Monroe, VA.

Physical Training Program

Trainees participated in 92 hours of physical training (road marches, calisthenics, running, rappelling, and swimming) and 41.5 hours of drill and ceremony during a 12-week stay at Parris Island. Of the time devoted to physical training, only the hours devoted to swimming and ceremony (19 hours) did not expose the trainees to the risk of lower extremity injury. Since 1976, the Marine Corps has allowed trainees to run in footwear other than boots, and about nine hours of training were spent running while wearing running shoes (recruit-purchased, not issued). Most of the running in shoes (80 per cent) occurred during the first five weeks of training. While the style and condition of the running shoes were not under control, the answers to questions about the condition of the shoes allowed some control for their effect.

Training was done by platoon unit—a group of approximately 65 persons who shared the same sleeping and eating facilities. During the first week, all trainees were required to take a diagnostic physical fitness test consisting of a one and one-half mile run, pull-ups, and situps. Passing a final physical fitness test consisting of a three-mile run, pull-ups, and situps was a requirement for graduation. Two long marches (seven miles and 10 miles) in boots and full combat equipment occurred in the 4th and 8th weeks, respectively.

Trial Design

Boots with polymer insoles were issued to trainees who were assigned to even-numbered platoons. Boots with a standard mesh insole were issued to members of odd numbered platoons. Exchanging or losing the insoles was minimized in two ways. First, virtually no interaction among trainees from different platoons was permitted, and all trainees within each platoon had the same insole type. Second, trainees were told at the beginning of training and periodically during training that they were to leave the insoles in their boots. Equipment inspections enforced compliance.

Ascertainment of Injury Endpoints

All Marine recruits who had lower extremity injuries were evaluated at a medical clinic on Parris Island. Injury information for this study was recorded on clinical data forms at the time of the clinic visit. Patients were transferred to Beaufort Naval Hospital for lower extremity x-rays.

Two radiologists independently evaluated the radiographs of injured recruits. The first radiologist, at Beaufort Naval Hospital, gave the primary radiologist's reading of the films. After the study was completed, the x-ray films were re-evaluated by another radiologist at Walter Reed Army Medical Center, Washington, DC. Both radiologists were blinded with respect to insole status and the second review was blinded concerning the primary evaluation.

Endpoint classification was based on functional impairment rather than solely on radiographic demonstration of a clear disruption of the integrity of bone cortex. Thirty-eight trainees had either x-ray-confirmed stress fractures or clinically serious stress reactions with x-ray evidence compatible with stress injury (cortical tunneling or lucency). Since recruits with pain in the lower extremities were managed conservatively, many stress reactions were prevented from progressing by repeated visits to the Parris Island medical clinic, which was open 24 hours a day. A number of individuals had only one film taken and, because of the delay in appearance of radiographic evidence following onset of symptoms, these early x-rays were frequently negative. Rather than drop from the analysis all persons with only early x-ray films, we pooled together x-ray-confirmed stress fractures with stress injuries having an incomplete set of radiographs but clinically significant symptoms and referred to them all as stress fractures.

The distribution of the 38 stress fractures was femur (5), tibia (14), fibula (5), calcaneous (1), and metatarsal (13). Most of these injuries were unilateral, but bilateral injuries occurred in the tibias of four recruits, and in the femurs, calcanei, or metatarsals of one recruit each.

Statistical Analyses

Fisher's exact test was employed to test the significance of proportions compared in two by two tables.[20] Where questionnaire responses followed an ordered sequence, a chi-square test for trend was employed.[21] All data were collected on mark-sense forms and edited and analyzed using version 4.0 of the Statistical Analysis System.[22]

Results

Table 1 presents fracture rates by age, ethnicity, and activity status. With only 37 recruits above age 25, the 21+ age group mainly represents the ages 21 to 25. Adjusted for physical activity, the relative risk for older vs younger recruits was 1.71. Whites, compared to other racial and ethnic groups, had a relative risk of 2.45. Intrinsic racial differences in susceptibility to stress fractures are given some support by these data.

Previously inactive trainees were at much greater risk, over 10 times the average of the other four activity categories. The inverse relation of risk to reported physical activity is strong in these data: the chi-square for linear trend was highly significant. Compared to trainees with above average activity levels, the relative risk of fracture for all others was 2.40 (95% C.I. = 4.58, 1.26).

As seen in Table 2, the fracture rate for the polymer insole group was slightly higher (1.35 per cent), but not very different from the rate in the standard insole group (1.13 per cent). Activity levels were similar in the two groups. In the less active subgroup and the more active subgroup, the fracture rates did not vary much by insole type.

We examined the fracture rates by location of stress injury (Table 3) to see whether the protective effect of increased shock absorption could be localized to the foot. We found no such effect.

TABLE 1—Stress Fracture Rates by Age, Ethnicity, and Activity

Variables	Fractures	Trainees	Rates (%)
Age (years)			
18–20	21	2074	1.01
21+	17	934	1.82
Summary* Relative Risk 18–20 vs 21+ = 1.71 (95% CI: 0.92, 3.21)			
Ethnicity			
White NH	32	2050	1.56
Hispanic	1	168	0.60
Black NH	5	743	0.67
Amerindian	0	20	—
Other/Unk	0	44	—
Summary* Relative Risk white vs all other = 2.45 (95% CI: 1.06, 5.68)			
Activity			
Inactive	3	25	12.00
Below Average	5	224	2.20
Average	15	924	1.62
Active	11	1197	0.91
Very Active	4	638	0.62

*Adjusted for history of reported physical activity (5 levels).

TABLE 2—Stress Fracture Rates by Boot Insole Type

	Total Group	Polymer	Mesh	Relative Risk	95% CI
N/PAR	38 /3025	21/1557	17/1468	1.17	(.62, 2.2)
Per Cent	(1.26%)	(1.35%)	(1.13%)		
Physical Activity = Average or Below	23 /1173	13/589	10 /584	1.29	(.57, 2.9)
	(1.96%)	(2.21%)	(1.71%)		
Physical Activity = Active or Very Active	15 /1835	8/960	7/875	1.04	(.38, 2.9)
	(0.82%)	(0.83%)	(0.80%)		

Since the cushioning and support from running shoes deteriorates with age, we examined whether the age of shoes was equivalent for polymer and standard insole groups. A modest trend of increasing fracture rates by age of shoes was observed (Table 4). The crude relative risk of 1.17 was unchanged after adjusting for age of shoes with two levels (less than one month, greater than one month). Cost of shoes in relation to fracture rates was examined. The rates varied only slightly among three categories of cost, with no apparent trend (Table 5).

Finally, we examined the differences for polymer vs mesh groups in a category comprising four other major injuries: plantar fasciitis, ankle sprains, knee strains and sprains, and achilles tendonitis (Table 6). Once again, no protective effect from the polymer insole was apparent.

Discussion

Published tests of the efficacy of shock absorbent materials in footwear have been few. On the other hand, many citations can be found which recommend the use of shock absorbing shoes or shock absorbent insoles based on a review of clinical cases.[14–17] Our review of the literature uncovered only one prospective study of the efficacy of shock absorbing insoles in preventing any type of overuse injuries. A 1985 study of 90 US Coast Guard recruits reports cushioned insoles successful in reducing "shock impact and shearing injuries"[23]; however, the impact and shearing injuries referred to were foot bruises, calluses, and blisters, not stress fractures or stress reactions. An unpublished 1986 report on the frequency of lower extremity injuries among 555 female Army basic trainees found no difference in stress fracture or stress reaction rates between trainees wearing a Sorbothane insole compared to those wearing a standard mesh insole.**

TABLE 3—Injury Rates by Insole Type Stratified by Injury Location

Location	Injury Type	Insole Type: Polymer	Insole Type: Mesh
Above the Foot	Stress Fractures	10 /1557 (0.64%)	14 /1468 (0.95%)
	Other Injuries	233 /1557 (15.0%)	253 /1468 (17.2%)
Foot	Stress Fractures	11 /1557 (0.71%)	3 /1468 (0.20%)
	Other Injuries	47 /1557 (3.01%)	38 /1468 (2.59%)

While observations on injuries in military training are not always generalizable to the general population, injury types in this and other military populations resemble civilian population reports. In basic training populations in the Army, five injury types are repeatedly cited as accounting for over 50 per cent of all training injuries: stress fractures, overuse injuries of the knee, plantar fasciitis, achilles tendonitis, and ankle sprains.[6,12,24] These same five injury types are cited as accounting for over 50 per cent of all injuries in civilian running and jogging programs.[14,16,25]

The uniformity of Marine training is advantageous for determining the effect of a simple training modification—the addition to the boot of a shock absorbing viscoelastic insole. Each drill instructor, platoon, and training company follows the same program during the 12-week training cycle. This rigid pattern is in contrast to the exercise patterns of runners presenting at an orthopedic or sports medicine clinic, where the entire spectrum of activity is represented.

The stresses of physical training were greater for those trainees who were initially physically unfit. Previous studies have also noted the association of fitness with injury risk, but they used data collected retrospectively, after the injury had occurred.[26] In contrast, the present study recorded responses to questions before training began. A simple multiple choice response to the question, "How would you describe your life before coming to Parris Island?", proved remarkably sensitive in discriminating trainees at risk of significant stress reactions. For those 25 categorized as "inactive" prior to their training, three (12 per cent) suffered a significant stress reaction. The other responses were associated with progressively lower rates of bone injury.

In civilian populations, the most frequently encountered abrupt change in physical activity is increased or newly adopted running, referred to as "training errors". These errors may be the most important cause of running injuries.[15,16] These clinical observations reflect the importance of what Powell, *et al*, have called "the stability of running habits."[27] Although we cannot give quantitative guidance from these data about the amount of activity change that leads to an injury risk, we can say that previously inactive recruits had a 10-fold excess risk of clinically significant stress reactions during basic training relative to the rest of the recruit population. The impact of the inactive group on the Marines (25 of 3,025) is small, but for the general

**Bensel C, Kaplan DB: Wear test of boot inserts. (Report). US Army Natick Research and Development Laboratories, Natick, MA, December 1986.

TABLE 4—Stress Fracture Rates by Age of Shoes

	<1 wk	1 wk–1 mo	1–6 mos	6 mos–1 year	>1 year	Total
N/PAR	20 /1792	7 /630	8 /422	3 /119	0 /44	38 /3007*
Per Cent affected	1.12	1.11	1.90	2.52	—	1.26

χ^2 trend = 2.54
One-tailed p = 0.056
*18 individuals had missing shoe information

population that proportion, and the injury consequences, would be much higher. The traditional recommendation to beginning runners (and the military's requirement for physically unfit trainees) to start slowly and to build up progressively the frequency, intensity, and duration of running, would seem well supported as a policy for preventing significant stress reactions of bone.

The distribution of stress reactions in this study (71 per cent of the total found in tibia or metatarsal) follows closely that observed in similar studies.[11,28] The authors of one study correlated their findings with the use of a heel snap to mark the cadence while marching. The heel snap has since been discontinued.

We observed a slight excess of clinically significant stress reactions of the metatarsals and calcanei in those recruits who used the Sorbothane insole. This excess was reflected, to a lesser extent, in higher overall injury rates below the ankle in the polymer insole group. The picture was reversed above the ankle, with a slight excess of fractures in the mesh insole group, and likewise an excess of all other types of injuries in the mesh insole group. Whether this pattern is causally related to the insole remains an intriguing question that needs testing.

These data also confirm earlier studies' reports of an excess of stress fractures or stress reactions in Whites compared to Blacks.[9] One explanation for this finding could be the lesser bone density and bone mass observed in Whites compared to Blacks.[29–32] If racial differences in bone density are important predictors, studies which can account for both bone density and physical activity history need to be done.

Of the 92 hours of physical training each recruit experienced, nine were spent in running shoes without an insole. Although we did not study running shoes in detail, the age of the shoes and the price paid provided information about condition of the shoes. Price, a correlate of overall quality, perhaps of shock absorbency, did not affect fracture rates, but increasing age of shoes did. The age of the shoe may indicate the degree to which its shock absorbent material has been compacted. On the other hand, the mechanical support provided by a shoe is also adversely affected by age. Without more data, cushioning effects cannot be separated from support effects. We also cannot assess the efficacy of shock absorbing insoles for moderate vs extreme amount of running.

Our findings suggest that the relation between footwear and lower extremity injuries is complex, that the simple model of the lower extremity as a projectile impacting upon shock absorbent material is not dynamic enough, and that shock or stress absorbency derives from the physiology and biomechanics of the lower extremities as well as from the physics of footgear. The most important modifiable factor in injury prevention was physical conditioning of the lower extremities.

TABLE 5—Stress Fracture Rates by Price of Running Shoes

	<$25	$25–40	$40+
N/PAR	19 /1555	12 /748	7 /722
Per cent affected	1.22	1.60	0.97

χ^2 trend = 0.5
p = n.s.

TABLE 6—Other Serious Lower Extremity Injuries by Insole Type

	Injuries	Trainees	Rate (%)
Polymer	101	1456	6.40
Mesh	97	1371	6.61

Relative Risk polymer vs. mesh = 0.98 (95% CI: 0.73, 1.31)

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Abstracts Sought for 1989 Public Health Conference on Records and Statistics

The 22nd biennial Public Health Conference on Records and Statistics will be held in Washington, DC, July 17–19, 1989. The conference is sponsored by the National Center for Health Statistics and serves as a major national forum for the latest advances in public health statistics.

The theme of the 1989 conference is "Challenges for Public Health Statistics in the 1990s." The conference offers an opportunity to assess the demands for health statistics in the next decade and how those demands can be met. Three major areas will be addressed:

- Promotion and Prevention
- Surveillance and Epidemiology
- Targeting Services

Presentations will be organized in sessions focusing on policies, programs and services; methodological and analytical issues; and trends, projections and goals. The conference will highlight topics of current health importance such as HIV (human immunodeficiency virus) Infection, Aging, Injuries, Maternal and Infant Health, Chronic Diseases, Preventable Morbidity and Mortality, Environmental Issues, and Ecological Analyses, as well as a special session on Marriage and Divorce Statistics.

Both invited and contributed papers will be featured. To be considered for presentation, NCHS must receive an abstract of a contributed paper by January 13, 1989. For more information on the call for papers, instructions on submitting an abstract for consideration, or to receive a conference invitation, contact Nancy Hamilton, OPEP, PHCRS, National Center for Health Statistics, 3700 East-West Highway, Room 2-12, Hyattsville, MD 20782. Tel: (301) 436-7122.

posttest counseling in this study, diverse forms of information and social support were given by the nurses who maintained contact with the participants. Testing was voluntary, with informed consent and under the strict condition of confidentiality. Because ethical and practical considerations prohibit experimenting with the factors mentioned above, their individual contribution to the effect of testing cannot be distinguished.

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APPENDIX

The relative importance of a sexual technique was determined by the computation of so called deviation contrasts, or deviations from the grand means.[10–12] A practical description of this procedure is found in the SPSSX Users Guide,[12] pp 494–495. New variables were computed for each individual and each sexual technique on pretest, posttest, and follow-up, by subtracting from the number of partners on a specific technique, the mean of the number of partners with whom the remaining techniques were performed. An example may clarify this procedure: if an individual performs masturbation with two partners, orogenital intercourse with six partners, and anogenital intercourse with eight partners, the relative importance of masturbation is $2-(6+8)/2=-5$, of orogenital intercourse $6-(8+2)/2=1$, and of anogenital intercourse $8-(2+6)/2=4$. Twelve months later, if this individual practices masturbation with eight partners, orogenital intercourse with two partners, and anogenital intercourse with zero partners, the relative importance of masturbation has increased to $8-(2+0)/2=7$, of orogenital intercourse has decreased to $2-(8+0)/2=-2$, and of anogenital intercourse has decreased to $0-(8+2)/2=-5$. It is concluded that a shift has occurred in the pattern of practiced sexual techniques from orogenital and anogenital intercourse in the direction of masturbation. Sexual techniques included in this procedure were: oral, manual, orogenital, anogenital, oroanal, anobrachial, anodigital, dildo use, rectal douching, scat and watersports.

Changes in the Distribution of Physicians in Rural Areas of Minnesota, 1965-85

TERRY DENNIS, MD

Abstract: The purpose of this study was to assess changes in the physician-population ratio in non-SMSA Minnesota counties between 1965 and 1985 using county specific data published by the American Medical Association. The physician-population ratio actually decreased by 2 per cent for primary care physicians and by 11 per cent for family practitioners in the non-SMSA counties. The large increase in the number of physicians in Minnesota has not translated into improved access to primary care physicians in Minnesota's rural areas. (*Am J Public Health* 1988; 78:1577–1579.)

Introduction

Physician manpower policy in the late 1960s was dominated by a shortage of primary care physicians in rural and inner-city areas. Federal and state governments responded to this problem by funding massive increases in the number of medical school graduates, the establishment of new primary care residency programs, and development of primary care service programs for the underserved.[1] Events in Minnesota which have mirrored the national response over the last 20 years resulted in a 50 per cent increase in the number of physicians in Minnesota between 1965 and 1977.[2]

By the late 1970s, concern had shifted from a physician shortage to an impending physician surplus.[3] It was asserted that a surplus of physicians would induce enough underutilized physicians to "diffuse" from the larger cities into ever smaller communities and resolve the physician manpower problem.[4,5]

This study was designed to assess whether "diffusion" was in fact occurring in Minnesota and, if so, whether it was sufficient to redress the deficits in primary care physicians identified 20 years ago.

Method

Population figures for each county in Minnesota were obtained from the 1970 census and projections of 1985 population.[6] Counties were classified as either "SMSA" or "non-SMSA" based on the 1977 US Department of Commerce definition of Standard Metropolitan Statistical Areas. The "non-SMSA" counties were then divided into the

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following groups using their 1985 population: less than 10,000; 10,000 to 24,999; 25,000 to 60,000 population.

Data on the distribution of non-federal Minnesota physicians by specialty group and county were obtained for the years 1965, 1975, 1980, and 1985 from American Medical Association publications.[7–10] The AMA data were categorized into four specialty groups:

- *Family Practice* which included family and general practitioners;
- *Medical Specialties* which included internal medicine, pediatrics, and their subspecialties;
- *Surgical Specialties* which includes general surgery, neurological surgery, obstetrics and gynecology, plastic surgery, colon and rectal surgery, thoracic surgery, and urology;
- *Other* which consists of all other specialties, hospital-based and unclassified physicians;
- Physicians classified as "Inactive" or "Other Professional Activity" were excluded.

The Family Practice and Medical Specialties groups were also combined into a single "Primary Care" group. This excludes Obstetrics-Gynecology and other surgical specialists as primary care providers.

The trend and change from 1965 to 1985 in the physician-population ratio for each specialty and county group were tabulated to assess if net "diffusion" was occurring.

Results

Table 1 illustrates the physician-population ratio in Minnesota by specialty group and category of county during the period 1965-85. The non-SMSA counties exhibited a marked decrease in family physician-population ratio from 1965 to 1975. Since 1975, replacement physicians have moved into these communities but the ratio remains higher than in 1965. In contrast, the family physician-population ratio in SMSA areas increased by 11 per cent during the two decades.

In contrast to family practitioners, medical, surgical and other specialists have increased in all communities. If internists, pediatricians, and family physicians are combined into a "Primary Care" classification, it is evident that the medium size counties (25,000 to 60,000 population) have experienced a slight increase in the primary care physician-population ratio (Table 2) while the smaller counties have experienced a decrease. The SMSA counties have had a major increase in this ratio although many of the physicians may be subspecialists in internal medicine and pediatrics. Despite the increase in non-primary care specialists, the total physician-population ratio has decreased in the counties of less than 25,000 population.

Table 3 summarizes these findings for all non-SMSA counties in Minnesota. There were fewer "Family Practice" physicians per population in 1985. Even when "Medical Specialties" are included with family physicians, there has been essentially no change in "Primary Care" physicians per population over the 20-year period. There are obviously more surgeons and other specialists available than in 1965, and an increase of 15 per cent in the overall physician-population ratio has occurred.

TABLE 1—Physician/Population (per 100,000) Ratio by County Population, Minnesota 1965–85

Practice Group/Year	Non-SMSA <10,000	Non-SMSA 10,000 to 24,999	Non-SMSA 25,000 to 60,000	SMSA
Family Practice				
1965	48.49	48.26	43.90	28.74
1975	38.79	38.10	35.60	25.17
1980	40.20	39.50	35.42	26.76
1985	43.97	39.50	38.71	32.45
Medical Specialties				
1965	2.42	2.15	7.90	23.34
1975	1.21	2.34	10.84	34.10
1980	3.77	3.62	13.84	42.46
1985	3.77	4.17	16.07	52.39
Surgical Specialties				
1965	4.85	4.49	13.52	29.27
1975	2.42	5.47	19.27	35.52
1980	2.51	5.98	19.00	37.97
1985	1.26	6.16	22.05	42.78
Other				
1965	0	2.74	10.97	74.40
1975	7.28	8.40	18.34	110.21
1980	7.54	8.52	19.00	110.86
1985	5.03	7.25	18.76	117.10
Total				
1965	55.76	57.64	76.29	155.66
1975	49.70	54.32	84.05	204.98
1980	54.02	57.62	87.26	218.01
1985	54.02	57.08	95.59	244.77

Discussion

The information presented here demonstrates that despite a greater than 60 per cent increase in the number of physicians, the goal of increasing the primary care physician-population ratio in Minnesota's smaller communities has not been achieved.

In contrast to the inadequate "diffusion" of primary care providers, other specialists are rapidly increasing in smaller communities. These specialists have undoubtedly brought new medical services to outstate Minnesota and taken over certain duties in surgery, orthopedics, and obstetrics from the core practice of family physicians. The amount and quality of primary care these surgical specialists are attempting to deliver is not known.

The potential relief offered the primary care specialists by the presence of their surgical colleagues must be considered in light of the changes within the primary care specialties themselves. The demographic and specialty characteristics of primary care providers have dramatically changed during the last 20 years. Some evidence[11–14] suggests that these changes have decreased the potential capacity of each physician.

This study did not include osteopaths. However, review of directories of the American Osteopathic Association[15,16]

TABLE 2—Primary Care/Population (per 100,000) Ratio by County Size, Minnesota 1965–85

Years	Non-SMSA <10,000	Non-SMSA 10,000 to 24,999	Non-SMSA 25,000 to 60,000	SMSA
1965	51	50	52	52
1975	40	40	46	59
1980	44	43	49	69
1985	48	44	55	85
% Change 1965–85	−6	−13	6	63

NOTE: Primary Care Physicians = Family Practice plus Medical Specialties (see Methods).

TABLE 3—Physician/Population per (100,000) Ratio in Non-SMSA Counties, Minnesota 1965–85

Practice Group	1965	1985	% Change
Family Practice	44	39	−11
Medical Specialties	5	11	114
Surgical Specialties	9	14	58
Other	7	12	81
Total	65	76	18
Primary Care	51	50	−2

reveals that 13 more osteopathic physicians were in non-metropolitan Minnesota in 1985 than in 1968. Seven of the new osteopaths were in counties of less than 10,000 population and would result in a small (6 per cent) increase in the physician-population ratio in the smallest counties. Aside from these smallest counties, the additional osteopaths would have had a negligible impact on the overall situation in non-metropolitan Minnesota.

The trends identified in this study may or may not reflect those in other areas of the US. Future investigations should also identify the impact of the changing demographic profile of physicians on primary care availability in smaller communities and the extent to which physicians classified as non-primary care specialists actually provide primary care, when they are located in smaller communities.

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The Accuracy of Industry Data from Death Certificates for Workplace Homicide Victims

HAROLD DAVIS, MD

Abstract: This study compared death certificate data on usual industry for workplace homicide victims in five urban Texas counties, with medical examiners' data on the industries where victims were working when injured. The overall positive predictive value of the death certificate data was 72 per cent. Death certificate data on usual industry underestimated the number of victims working in high-risk industries when injured, partly because of victims whose usual industry was recorded as student, housewife, or military personnel. (*Am J Public Health* 1988; 78:1579–1581.)

Introduction

To determine the industry of a worker who was a homicide victim, recent studies[1–3] have used death certificate data on usual industry (the industry in which a person worked during most of his/her working life). However, at the time of injury, a workplace homicide victim might not have been working in his/her usual industry (e.g., educational services) but rather in a second, different one (e.g., gasoline service stations). Therefore, death certificate data on usual industry might not accurately document the victim's industry when injured. Although studies have examined the accuracy with which death certificate data document a person's usual industry,[4–7] none have examined the accuracy with which these data document a workplace homicide victim's industry at the time of injury.

In most urban Texas counties, medical examiners investigate all injury-related deaths. For workplace homicides, the medical examiners' records usually contain information on the victim's industry when injured. Using medical examiners' data from five urban Texas counties as the standard, I studied the extent to which death certificate data on usual industry accurately document the industries where victims were working when injured.

Methods

For the deaths they investigate, medical examiners in Texas complete certain parts of death certificates, including those on cause of death; accident, suicide, or homicide; injury at work; how a fatal injury occurred; and place of fatal

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injury (e.g., bar, street). Funeral directors also complete parts of these death certificates, including that on usual business or industry.

Cases of workplace homicide were identified via two steps. First, a review of Texas death certificates identified the death of any person who was ≥16 years old; whose cause of death was coded by the International Classification of Diseases (8th[8] and 9th[9] revisions) to "homicide" (ICD Codes E960-969); who was injured in Texas; whose death certificate had a positive response to "injury at work?"; and who died in the period 1975 through 1984 in Harris, Dallas, Tarrant, Bexar, or Travis Counties (which contain the cities of Houston, Dallas, Fort Worth, San Antonio, and Austin, respectively). In the second step, medical examiners' records on these deaths were reviewed to confirm that they were in fact homicides that had occurred in the workplace.

Data from death certificates on usual industry and from medical examiners' records on industry when injured were assigned three-digit codes according to the 1980 census industry classification system.[10] The coding system has been described previously.[2] A death certificate entry was considered accurate if its three-digit industry code was identical to that for the entry on the medical examiner's record. Only cases with industry data on both the death certificates and medical examiners' records were included.

Results

Of the 533 confirmed cases included in the analysis, 386 had death certificate industry data that were accurate, yielding a positive predictive value of 72 per cent (Table 1). The positive predictive value of death certificate data on male victims (320 of 446 industries accurate, 72 per cent, 95% CI 67–76 per cent) was similar to that for female victims (66 of 87 industries accurate, 76 per cent, 95% CI 66–84 per cent).

Of the industries previously shown to have high risks of workplace homicide (Table 1), six of seven had death certificate usual-industry data with positive predictive values for industry when injured of ≥89 per cent. In contrast, the death certificate usual-industry entries "student", "military personnel", and "housewife" never correctly described the victims' industries when injured. For the 28 victims with these usual-industry entries on their death certificates, medical examiners' records revealed that when injured they were working in food-bakery-and-dairy stores (seven persons), gasoline service stations (seven), eating-and-drinking places (five), or other industries (nine).

For five of seven high-risk industries, death certificate usual-industry data underestimated the number of victims working in them when injured (Table 1). This was most marked for eating-and-drinking places, which had 96 persons listed on medical examiners' records but only 61 persons listed with that usual industry on death certificates.

Of the 147 death certificates whose usual-industry data did not reflect the industry when injured, 60 (41 per cent) had death certificate description-of-injury or place-of-injury portions that correctly described the victims' industries when injured. For the other 87 death certificates, these portions either had no description of the victims' industries (86 of 147 death certificates, 59 per cent) or an incorrect description (one of 147, <1 per cent).

TABLE 1—Accuracy of Death Certificate Industry Data for Workplace Homicide Victims, 16 Years old or older, 1975–1984, in Five Urban Texas Counties[a]

Industry	Medical Examiners' Records (industry when injured)	Death Certificates (usual industry)		
		Listed	Accurate[b]	PPV[c] (95% CI)
High-risk Industries				
Food-bakery-and-dairy stores	114	99	92	93 (86, 97)
Eating-and-drinking places	96	61	59	97 (89, 99)
Gasoline service stations	49	38	37	97 (87, 100)
Justice, public order, and safety	33	35	31	89 (74, 95)
Taxicab service	25	22	22	100 (85, 100)
Detective and protective services	9	10	7	70 (40, 89)
Hotels, motels, and lodging places	8	7	7	100 (65, 100)
All high-risk industries	334	272	255	93 (90, 96)
Other Industries				
Students	0	19	0	0 (0, 17)
Military	0	6	0	0 (0, 39)
Housewives	0	3	0	0 (0, 56)
Other	199	233	131	56 (50, 62)
All Industries	533	533	386	72 (68, 76)

[a]Harris, Dallas, Tarrant, Bexar, and Travis Counties.
[b]Accurate indicates that death certificate data on usual industry correctly described a decedent's industry when injured.
[c]PPV = positive predictive value (95% confidence intervals).

Discussion

This study found that when the usual industry entered on a workplace homicide victim's death certificate is a high-risk one, the victim very likely was working in that industry when injured. In contrast, when the usual industry entered on the death certificate is a low-risk one (and especially if the entry is student, housewife, or military personnel), the victim often was not working in that industry when injured but in a high-risk one. Due to the disparities between the usual-industry entries on death certificates and the industries where victims were actually working when injured, death certificate data on usual industry tend to underestimate the number of victims working in high-risk industries when injured[4].

Death certificate data on usual industry can correctly identify which industries have high risks of workplace homicide, but estimation of these risks might be made more accurate by supplementing the usual-industry data with data from the description-of-injury and place-of-injury portions of death certificates. Nevertheless, in 59 per cent of deaths with inaccurate usual-industry data, these latter two portions do not describe the victims' industries at the times of injury. Persons completing description-of-injury and place-of-injury portions of death certificates can increase the usefulness of death certificates for the study of workplace homicides by including in these portions descriptions of the industries where victims were working when injured.

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Nurse Administration of Sleep Medication: A Comparison of Registered Nurses and Licensed Practical Nurses

PAMELA A. ROBERS, PHD, TERRI SMITH MOORE, MBA, AND BONNIE L. SVARSTAD, PHD

Abstract: This study examines differences in the administration of PRN (pro re nata or give "as needed") sleep medication between licensed practical nurses and registered nurses working in long-term care facilities. Work characteristics, characteristics of the residents cared for, and the extent of orders and administration of PRN sleep medications were similar in the two groups. No significant differences in the administration of PRN sleep medication were found. (*Am J Public Health* 1988; 78:1581–1583.)

Introduction

Administration of medication in long-term care facilities (LTCFs) is a function performed primarily by licensed practical nurses (LPNs) and registered nurses (RNs). According to the National League for Nursing, "LPNs/LVNs are prepared to function under the guidance of a registered nurse or licensed physician. . ."[1] However, one report indicates that even in the structured setting of the hospital, nine out of ten RNs responded that, in their facility, LPNs routinely administered medications without the direct supervision of an RN.[2] In the less structured nursing home setting, the autonomy of LPNs is likely to be even greater than in hospitals.[3–6]

When medications are prescribed PRN (pro re nata or give "as needed") the ultimate decision is delegated to the nurse in charge of administering medication. Studies examining PRN prescribing in LTCFs report that residents have an average of three to four orders for PRN medications during a time frame of 30 days or less.[7–16] Yet little is known about the factors affecting the nurse's medication decisions.

Sleep medications are commonly ordered for nursing home residents,[8,13,16,17–19] and about 50 per cent of sleep medication orders are written for PRN administration[13,20]; elderly individuals are particularly sensitive to and more likely to experience adverse effects than would younger individuals.[21–30] We examined the differences in sleep medication administration patterns between RNs and LPNs working in skilled nursing facilities.

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Hunting Firearm Injuries, North Carolina

THOMAS B. COLE, MD, AND MICHAEL J. PATETTA, MA

Abstract: To determine the percentage of unintentional firearm-related injuries associated with hunting and to identify risk factors, we conducted a retrospective, descriptive survey of all hunting firearm injuries identified by two North Carolina surveillance systems. Almost one-third of unintentional shooting deaths are hunting-related, and young hunters appear to be at greatest risk of injury. Safety instruction and wearing highly visible clothing should be encouraged; controlled studies should test the effectiveness of these preventive measures. (*Am J Public Health* 1988; 78:1585–1586.)

Introduction

Unintentional firearm injuries are a small but significant fraction of all firearm injuries in the United States. Approximately one-fifth of unintentional firearm injuries have been attributed to hunting.[1,2] We retrospectively reviewed all reports of fatal and nonfatal hunting firearm injuries identified by two North Carolina surveillance systems for the two-year period 1984–85 and all fatal hunting firearm injuries for the five-year period 1981–85.

Methods

Data from two sources, Outdoors Accident (OA) Reports of fatal and nonfatal hunting injuries and Medical Examiner (ME) reports of fatal injuries for the years 1984 and 1985, were merged. OA Reports were obtained from the North Carolina Wildlife Resources Commission and ME Reports from the North Carolina Office of the Chief Medical Examiner. ME Reports are required by law to be submitted for all fatal injuries,[1] and are therefore assumed to be population-based. Because ME Reports are not computer coded according to whether the injury was hunting-related, all unintentional firearm deaths were reviewed as well as all intentional firearm deaths with codings that suggested an increased probability of being hunting-related (outdoor non-residential location, non-handgun weapon, or undetermined manner).

A case was defined as a firearm injury for which the shooter was known to be hunting. The period of hunting, as defined on the Outdoors Accident Report form, extends from the time when "hunting equipment is removed from its natural place of storage until it is returned to storage after the hunt."

The number of hunting licenses issued in North Carolina in 1985 and the number of deer reported killed in 1985 (using data obtained from records of big game licenses) were obtained from the North Carolina Wildlife Resources Commission. US Census data from 1980 were used to assign a designation of rural or urban to each county.

All statistical tests are by Fisher exact test, two-tailed. Relative risks are reported with exact 95% confidence intervals.

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Results

There were 78 fatal and nonfatal hunting firearm injuries in 1984–85. The ratio of injury to hunters was 11.3 per 100,000 hunting licenses sold per year. The ratio of firearm injuries to deer hunters during this period was 34.7 per 100,000 deer killed per year. The rate of hunting firearm injuries for North Carolina residents was 0.6 per 100,000 persons per year.

Of the 78 hunting firearm injuries reported during 1984–85, 22 (28 per cent) were fatal, 30 per cent of all unintentional firearm fatalities in North Carolina during 1984–85. Fatal hunting injuries might account for less than 28 per cent of the total because nonfatal firearm hunting injuries are probably underreported to the Wildlife Resources Commission.

Of all 78 victims injured in 1984–85, 72 (92 per cent) were hunting at the time of injury, and 85 per cent of the injuries occurred during deer hunting season from mid-October to January 1. Most other game can be legally hunted only during the fall and early winter as well. Hunting firearm injuries were reported from 47 of the 100 counties in North Carolina.

Of the 66 shooters whose counties of residence were known, 46 (70 per cent) were hunting in their home county, and 14 (21 per cent) were hunting in a county that bordered on their home county.

The median age of victims injured by hunting firearms in the state in 1984–85 was 25 years. The median age of 68 shooters was 23 years. Of the 10 shooters <15 years of age, five (50 per cent) were hunting without the supervision of an adult (≥21 years old). Of the 12 shooters 15–17 years of age six (50 per cent) were hunting without adult supervision. Seventy-three of 78 victims (94 per cent) and 70 of 70 (100 per cent) shooters whose sex was recorded were male.

All hunting-related injuries reported in 1984–85 were unintentional. The relationship of shooters to victims is described in Table 1. In only 18 cases was the shooter a stranger.

Only 8/67 victims were wearing highly visible (red or orange) clothing when they were shot. Overall, 25/78 (32 per cent) victims were shot under conditions of good visibility, in daylight when the weather was clear in an open area (defined as the absence of woods, underbrush, or tall weeds, crops, or grass between victim and shooter) on level ground.

A narrative description of events was provided for 74/78 cases. From these narratives, 165 unsafe hunting practices were determined to have been committed by shooter or victim or both, for a mean of 2.2 unsafe hunting practices per injury (Table 2). These 165 practices included 22 different types of unsafe hunting practice. In every hunting firearm injury case for

TABLE 1—Relationship of Shooters to 78 Victims by Hunting Party, North Carolina, 1984–85

Relationship of Shooter to Victim	Victim and Shooter in Same Hunting Party	Victim and Shooter not in Same Hunting Party
Family member	8	2
Friend	32	1
Self	17	—
Stranger	0	18
Total	57	21

TABLE 2—Number and Percentage of the Five Most Common Types of Unsafe Hunting Practice Reported for 78 Hunting Firearm Injuries, North Carolina, 1984–85

Unsafe Hunting Practices	Number (Percent) (N=103)
Firearm discharged although trigger was said not to be pulled*	32 (31)
Carrying firearm incorrectly**	20 (19)
Victim covered by shooter swinging on game	19 (18)
Victim out of sight of shooter	17 (17)
Victim mistaken for game	15 (15)

*By 8–19-year-olds 17/26 vs ≥ age 20, 15/42. Relative risk 1.8 (95% CI 1.1, 3.0)
**By 8–19-year-olds 14/26 vs ≥ age 20, 6/42. Relative risk 3.8 (95% CI 1.6, 9.8)

which a narrative was provided, there was at least one unsafe hunting practice. Two unsafe hunting practices were more likely to be committed by younger shooters. Fourteen of 61 shooters reported having completed a hunter safety course.

Review of Medical Examiner reports for the five-year period 1981–85 found that blood alcohol tests were performed for 41 of 50 fatal hunting firearm victims. Of these 41, six were positive. Among the 14 victims of self-inflicted injury who were tested for alcohol, five had positive blood alcohol tests, two of which were suicides.

Discussion

Estimates of the percentage of unintentional firearm injuries attributable to hunting in the United States vary.[1–3] Discrepancies could be due to the method used to classify unintentional shootings as being hunting-related. We believe our case-by-case review to be the best approximation of the percentage of unintentional firearm fatalities associated with hunting which has been published to date.

Most shooters and victims in this study were young, and half of the child and adolescent shooters were hunting without adult supervision. By contrast, 60 per cent of victims of fatal hunting firearm injury in a Swedish study were 40 years of age or older.[4] The probable explanation for this difference in age is that Swedish children do not hunt; in Sweden only airguns can be legally owned or used by persons under the age of 18.

North Carolina law prohibits the use of firearms by children <12 years old unless they are accompanied by an adult. In this series of hunting firearm injuries, only two shooters were <12 years old: one was supervised by an adult and the other was not. However, 26 shooters, half of whom were unsupervised, were <20 years of age. Adult supervision of child and adolescent hunters, particularly those who have not had formal instruction in the safe use of firearms, might prevent many hunting firearm injuries.

Thirty-nine states including North Carolina require hunters to wear fluorescent orange clothing while hunting to increase their visibility to other hunters; fluorescent orange clothing was not required for hunters during the periods of this study, however. Clothing of this color has been shown to be more visible to the human eye, particularly in low-light conditions, than clothing of other colors.[5] Without knowing how many hunters who were not injured by firearms in 1984–85 were wearing fluorescent orange clothing, we could not determine whether wearing fluorescent orange was statistically associated with a decreased risk of hunting firearm injury.

ACKNOWLEDGMENTS

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Smokeless Tobacco Use and Attitudes toward Smokeless Tobacco among Native Americans and Other Adolescents in the Northwest

ROBERTA L. HALL, PHD, AND DON DEXTER, DMD

Abstract: A survey of 1,180 sixth, ninth, and eleventh graders in three school districts in the State of Washington found that 34 per cent of male Native Americans, 24 per cent of female Native Americans, 20 per cent of male non-natives and 4 per cent of female non-natives are current users of smokeless tobacco products. In all areas and groups, the best predictor of whether an adolescent is a user is the use pattern of friends. (*Am J Public Health* 1988; 78:1586–1588.)

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Introduction

Smokeless tobacco in the form of chewing tobacco and snuff is a highly addictive substance that poses many health hazards for the growing number of children and adolescents who use it.[1–4] Nicotine and other harmful chemicals in tobacco are absorbed into the user's body through the mucosa. Oral leukoplakia, a white thickening of the oral mucosa, is a common consequence of use and is considered a precancer indicator.

Many studies have reported a growing use of smokeless tobacco in the general population[5,6] but only a few have focused on Native Americans. Schlife[7] in a study of 2,454 girls and 2,511 native and non-native boys in rural Alaska ages 5 to 18 found that 28 per cent of the girls and 34 per cent of the boys used smokeless tobacco products. In a study of 144 Native Americans, grades kindergarten through 12 from Alaska and Washington, Schinke, *et al*,[8] found that 19 per

cent of the females and 42 per cent of the males had used smokeless tobacco at least 20 times.

Methods

To investigate risk factors predictive of smokeless tobacco in native populations and to determine attitudes that suggest mechanisms for intervention, the Northwest Portland Area Indian Health Board contracted with staff at Oregon State University for a survey of school-aged youth. Three Washington State school districts where Native American youth comprise a significant minority were chosen to represent the geographic diversity of the Northwest. Neah Bay is a small, isolated community at the extreme northwestern tip of the state; Colville is a large, isolated reservation in the center of the state; and the third district lies adjacent to the Lummi Reservation near Bellingham in the urbanized western sector. In 14 schools in these areas, all sixth, ninth, and eleventh graders at school on the day of the survey in May 1987 were asked to participate. The refusal rate was 3 per cent and the absentee rate was 13 per cent. A total of 1,180 took part.

The survey ascertained respondents' use, individual traits (such as age, sex, and tribe), use practices of family and friends, and attitudes. Respondents who said they had never used smokeless tobacco were defined as non-users (640 or 54 per cent); those who currently use smokeless products constituted the users category (197 or 17 per cent); those who reported they had used but no longer used the product were categorized as former users (343 or 29 per cent). Comparisons were made between use categories and age, sex, Indian status, grade, area, attitudes, and use by friends and relatives. Multiple regression models were developed to test predictive models of use status and frequency of use.

Results

Use of smokeless tobacco products varies by ethnic status, sex, and, to a lesser extent, by grade level (Table 1), but these factors interact. Native American use in all areas is higher than non-native use. The contrast is particularly striking among females: whereas only 4 per cent of the non-native female sample use smokeless products, 24 per cent of Native American females are users.

The Native American sample is characterized by a younger age at first use; 33 per cent of native former users and 57 per cent of users tried smokeless tobacco before the age of 10 while only 17 per cent of non-native former users and 24 per cent of users had done so.* A larger percentage of non-users (80 per cent) had never smoked cigarettes than former users (49 per cent) or users (42 per cent).

Multiple regression was used to test a model of variables predicting whether a person is a user, former user, or non-user of smokeless tobacco with separate tests for males and females. Variables included age, grade in school, Native American or non-native ancestry, proportion of friends who use smokeless tobacco, smoking of cigarettes, and use by siblings, fathers, mothers, and other relatives. For boys, the model produced an adjusted R^2 of 0.4593. The number of friends who use the product contributed the greatest portion of the variance explained; other explanatory variables included: having smoked cigarettes, use by siblings, use by other relatives, and use by the father (see Table 2). In females the most important variable also was friends' use but include native ethnicity, grade in school (negatively related to use of smokeless tobacco), and age (positively related) in addition to variables found in the male model. The adjusted R^2 is 0.3935.

Regression models with the same variables were developed to predict the frequency and quantity of use but were less successful, in part due to smaller sample size. The only important variables for boys were friends' use and age at first use (adjusted R^2=0.2329); for girls it was friends' use (adjusted R^2=0.1619).

Attitudes toward smokeless tobacco vary significantly among users, former users, and non-users, but not at all by sex, ethnic group, area, or age. Reasons given for starting to use and continuing to use smokeless tobacco related to pleasure and taste with fewer than 10 per cent indicating peer pressure or addiction. Principal reasons non-users gave for not trying it focused on health issues and the unpleasant smell and appearance of smokeless products.*

Discussion

Although the rate of use of smokeless tobacco varies greatly by area, sex, and ethnic status, the pattern of attitudes held by users cross-cuts other differences. Included in the

TABLE 1—Use of Smokeless Tobacco, in Percentages, by Area, Ethnic Group, Grade, and Sex

Category	(N)	Category of Use: Non-User	Former User	Current User
Grade Level				
6	(379)	69	21	10
9	(428)	45	33	21
11	(373)	49	32	18
Area and Ethnic Group*				
Colville Native American	(128)	21	31	44
Colville non-native	(315)	57	34	9
Bellingham Native American	(92)	47	33	20
Bellingham non-native	(529)	63	23	14
Neah Bay Native American	(37)	46	40	14
Neah Bay non-native	(9)	33	56	11
Ethnic Group and Sex*				
Male Native American	(137)	29	37	34
Male non-native	(438)	46	34	20
Female Native American	(120)	45	31	24
Female non-native	(415)	76	20	4
Total Sample	(1180)	54	29	17

*Seventy respondents did not list their sex or ethnic group or both; of these, 41 per cent are non-users, 33 per cent are former users, and 26 per cent are users.

TABLE 2—Multiple Regression Model Predicting Use of Smokeless Tobacco

Variables	Boys: Regression Coefficient	Boys: Standard Error	Girls: Regression Coefficient	Girls: Standard Error
Proportion of friends using	.308	.029	.298	.035
Use of cigarettes	.473	.058	.215	.050
Native ancestry	.112	.063	.242	.059
Use by siblings	.244	.076	.242	.068
Use by father	.239	.078	.164	.072
Use by other relatives	.241	.055	.124	.049
Age	−.002	.036	.108	.038
Grade in school	.039	.038	−.122	.039

*Data available on request to author.

pattern is a denial of the role of peer pressure even though friends' use is the most important predictor, and denial of addiction, even though in both sexes and ethnic groups and in all three areas over half of the users said they "would quit if they could." These data suggest that intervention programs need to work with issues of friendships, pleasure, appearance, and health, and suggest that some adolescent users would respond positively to cessation programs if they were available.

ACKNOWLEDGMENTS

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NIH Consensus Statement Available on Perioperative Red Cell Transfusion

Transfusion of red cells is a life-saving measure in the management of a variety of medical and surgical conditions. The AIDS epidemic has recently raised the level of apprehension regarding the transmission of infectious disease by blood transfusion. Furthermore, there is new information about the significance of anemia in the perioperative period. These developments have stimulated a reexamination of the benefit-to-risk relationship for transfusion therapy.

To assess this procedure, the National Institutes of Health held a Consensus Development Conference on Perioperative Red Cell Transfusion, June 27–29, 1988. Based on scientific data presented, a consensus panel, drawn from the medical profession, blood blanking organizations, and the general public wrote a consensus statement. Following is a brief suramary of the panel findings:

- A hemoglobin of less than 10 g/dL or a hematocrit of less than 30 per cent indicates a need for perioperative red cell transfusion.
- Most patients with hemoglobin values >10 g/dL rarely need perioperative transfusions, whereas those with acute anemia and hemoglobin values <7 2/dL will more frequently need blood.
- Deciding to transfuse red cells depends on clinical assessment aided by laboratory data such as: arterial oxygenation, mixed venous oxygen tension, cardiac output, the oxygen extraction ratio, and blood volume, when indicated.
- There is no evidence that mild to moderate anemia contributed to perioperative morbidity.
- The number of homologous transfusions should be kept to a minimum.
- Although homologous red cell transfusions are becoming safer, they should not be considered substitutes for good surgical and anesthetic techniques.
- A variety of alternatives to homologous transfusions now are available, including the use of autologous blood, collected perioperatively, and by intraoperative blood salvage, which appears to be safe in some applications.

Free, single copies of the complete *NIH Consensus Statement on Perioperative Red Cell Transfusion* may be obtained from the Office of Medical Applications of Research, Building 1, Room 216, National Institutes of Health, 9000 Rockville Pike, Bethesda, MD 20892.

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Potential for Exposure to Tick Bites in Recreational Parks in a Lyme Disease Endemic Area

RICHARD C. FALCO, PHD, AND DURLAND FISH, PHD

Abstract: Eight recreational parks located in a Lyme disease endemic area of southern New York State were surveyed for the presence of ticks during the summer of 1985 by drag sampling. *Ixodes dammini*, the primary vector of Lyme disease in the northeast, was found in all but one park and accounted for 580 (91.8 per cent) of the 632 ticks collected. Of these, 18 per cent were larvae, 80 per cent were nymphs, and 2 per cent were adults.

An *I. dammini* encounter distance, defined as the mean number of meters traveled before encountering a nymphal or adult *I. dammini* on a drag cloth, ranged from 36 m in high-risk parks, to infinity (no tick encounters). Generally, areas of high use presented higher encounter distances (lower risk) than those of the entire park. Two of the three parks with the highest annual attendance also had the highest *I. dammini* population indices as projected from our sampling regimen. These results indicate that recreational parks in Lyme disease endemic areas represent a substantial human risk for tick bites and Lyme disease. (*Am J Public Health* 1989; 79: 12–15.)

Introduction

Lyme disease is a potentially serious tick-borne spirochetosis with a world-wide distribution.[1] It is particularly prevalent in the northeastern United States, where the deer tick, *Ixodes dammini*, is the primary vector.[2,3] Westchester County, New York, located to the immediate north of New York City, is an endemic area for Lyme disease,[4] and has an abundance of recreational parks, campgrounds, and nature preserves.

Lyme disease can be difficult to diagnose,[5] and delayed treatment may lead to cardiac and neurologic abnormalities,[6,7] as well as chronic arthritis.[8] Because of this potential for serious sequelae, early recognition and treatment of Lyme disease is important. Preventing the infection is preferable, but this requires knowledge of tick infested areas in order that tick bites may be avoided. Also, because many Lyme disease patients do not recall receiving a tick bite,[9] knowledge of exposure to infested areas may be helpful in diagnosing the disease.

The presence of homophilic ticks in recreational areas can constitute a significant public health threat.[10] For example, 45 cases of tick-borne relapsing fever were reported to have been acquired at Grand Canyon National Park during a single outbreak in 1973.[11] A recreational area in Tennessee, where serologic evidence from mammals and pathogen isolations from ticks showed Rocky Mountain spotted fever and tularemia to be endemic,[12] was found to have an abundance of the vector species *Amblyomma americanum* and *Dermacentor variabilis*.[13]

Commonly used methods for sampling ticks include the use of carbon dioxide-baited tick traps,[14] examination of host animals,[15] and drag sampling.[16] Although recreational areas elsewhere have been sampled for other tick species,[11,13] there is no published regimen to efficiently sample recreational parks for the presence of *I. dammini* in Lyme disease endemic areas.

In order to identify recreational areas with potential for subjecting the public to risk of acquiring tick bites, and possibly Lyme disease, a project was begun in 1985 to design an efficient sampling regimen to assess the species and abundance of ticks of public health importance in parks operated by the Westchester County, New York, Department of Parks, Recreation, and Conservation.

Methods

Eight recreational parks operated by the Westchester County, New York, Department of Parks, Recreation, and Conservation were selected for this study. Selected parks attracted large numbers of summer visitors and/or were located near clusters of Lyme disease cases reported to the Westchester County Department of Health. Sampling was performed in 1985 during the months of June and July when the nymphal stage of *I. dammini*, which is responsible for the majority of Lyme disease cases, is most active.[4,17] These months are believed to represent the greatest period of public exposure to Lyme disease.

Because the nymphal activity peak spans a narrow range of time, an efficient sampling regimen was required. Drag sampling provides information for a greater land area per unit effort than does either animal trapping or carbon dioxide-baited tick traps. Therefore, all parks were surveyed for ticks by drag sampling. This method involved dragging a one square meter flannel cloth, fixed to a wooden handle, over substrate and vegetation through the parks.[16] The cloth was examined every 20 meters for the presence of ticks, which were collected and preserved in alcohol for later identification. Because ticks captured by this method were seeking hosts, they represent the portion of the resident tick population that presents a risk for humans.

Transects were established through each park in order to systematically sample ticks from all habitats. Each transect consisted of a straight line through the park extending from border to border. To avoid sampling bias caused by vegetation or topography, orientation of the transects was in north-south and/or east-west directions. Because sampling time was limited, the objective was to sample as close to 1 per cent of the park as possible without spending more than five man-days at any one park. The number of transects sampled at each park therefore varied from one to five, with a mean of 3.9 (SD=1.45). Total sampling time was recorded for all areas.

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In addition to the transects, areas of high public usage were designated by Westchester County Department of Parks, Recreation, and Conservation personnel to be included in the survey. As in the transects, these locations were sampled by drag sampling. High usage areas included wooded and grassy areas, as well as trails and developed land. Such areas were designated for all parks except one, which was not open to the public during the sampling period.

An "encounter distance" for *I. dammini* was determined for each park in order to compare relative risk among parks. Defined as the mean number of meters traveled before encountering a nymphal or adult *I. dammini* on the drag cloth, an encounter distance was calculated as the reciprocal of the mean number of ticks collected per linear distance sampled. This distance has an inverse relationship to risk since parks with lower encounter distances have a greater likelihood of a person encountering a tick. Encounter distances were based on nymphs and adults only, as larvae are not considered to be epidemiologically important vectors of Lyme disease.[18,19]

The potential public health hazard from exposure to *I. dammini* in each recreational park was estimated by comparing the 1984 park attendance figures provided by the Westchester County Department of Parks, Recreation, and Conservation, to an *I. dammini* population index. This index was based on the number of nymphal and adult *I. dammini* found in the areas sampled extrapolated to the total area of the park. Park attendance figures were not available for two of the parks sampled.

Results

A total of 632 ticks, representing five species, were collected during this study (Table 1). The most abundant species was *I. dammini* (91.8 per cent) and it was found in all parks but one. Of the 580 total *I. dammini* collected, 105 (18.1 per cent) were larvae, 463 (79.8 per cent) were nymphs, and 12 (2.1 per cent) were adults (nine males, and three females). *D. variabilis* was the second most common species (5.5 per cent), and was found in all but two parks. *I. dentatus* (1.6 per cent), *Haemophysalis leporispalustris* (0.9 per cent), and *I. cookei* (0.2 per cent) were also collected.

Sampling time totaled 173.3 man hours and ranged from 6 hrs 20 mins to 35 hrs 10 mins per park, with a mean of 21 hrs 40 mins (Table 2). This effort resulted in less than 1 per cent of the total land area for each park being sampled. Sampling time was dependent upon density of vegetation and number of immature ticks collected. A mean of 11 minutes was required per 100 meters sampled. However, sampling time for particular transects and high use areas ranged from two to 25 minutes per 100 meters.

TABLE 1—Tick Species Composition of Westchester County, NY, Parks in 1985

Park	Total Ticks	*Ixodes dammini*	*Dermacentor variabilis*	*Ixodes dentatus*	*Haemophysalis leporispalustris*	*Ixodes cookei*
A	78	69	9	0	0	0
B	27	3	14	4	6	0
C	189	185	4	0	0	0
D	2	0	0	2	0	0
E	136	135	1	0	0	0
F	42	40	0	2	0	0
G	12	11	1	0	0	0
H	146	137	6	2	0	1
Total	632	580	35	10	6	1

TABLE 2—Sampling Effort for 1985 Westchester County, NY, Park Survey

Park	Transect Distance (M)	High Use Area Distance (M)	Total Distance (M)	% of Park Sampled	Man Hours (Hrs:Mins)
A	2260	2920	5180	0.96	27:40
B	3760	6600	10,360	0.51	35:10
C	6040	0	6040	0.35	23:40
D	1720	2160	3880	0.64	11:00
E	2440	1560	4000	0.13	17:40
F	6060	1940	8000	0.85	23:10
G	3160	200	3360	0.89	6:20
H	4880	6320	11,200	0.07	28:40

Calculated *I. dammini* encounter distances for the total sampling area at each park ranged from 36 m to infinity, where no *I. dammini* was found (Table 3). A comparison of the encounter distances for the eight parks sampled in this study reveals two distinct groups: high-risk and low-risk parks. High-risk parks had relatively low encounter distances of 36 m, 38 m, 88 m, and 118 m (Table 3). Low-risk parks had encounter distances of 208 m, 303 m, 3,333 m and infinity. The lowest encounter distance in this group is almost double that of the highest in the high-risk group. The extremely high encounter distances of 3,333 m and infinity suggest that there is little or no chance of encountering *I. dammini* at these parks.

I. dammini encounter distances for the high use areas alone (Table 3) were, in most cases, higher than those determined for the entire parks. In only two parks were the encounter distances lower in the high use areas. Total area encounter distances were 3,333 m and 118 m at these parks, while encounter distances for the high use areas were 2,000 m and 112 m, respectively.

The total public health risk from ticks in these recreational parks, based upon their attendance figures and *I. dammini* population indices, is shown in Table 4. Of the three parks with the highest attendance figures, two also yielded the highest *I. dammini* population indices. This suggests a greater potential for public exposure to tick bites at these parks.

Discussion

The large number of *I. dammini* collected suggests that it is the most abundant tick species encountered by humans

TABLE 3—Encounter Distances (M) for Nymphal and Adult *I. Dammini* in Westchester County, NY, Parks in 1985

Park	All Areas	High Use Areas
A	88	172
B	3333	2000
C	38	No areas
D	No encounters	No encounters
E	36	313
F	208	2000
G	303	No encounters
H	118	112

TABLE 4—Potential Human Exposure to *I. Dammini* in Westchester County, NY, Parks in 1985

Park	*I. dammini* Population Index	1984 Attendance
A	6151	11,477
B	584	84,606
C	44,812	Not available
D	0	22,489
E	87,024	44,675
F	4484	Not available
G	1231	1,784
H	146,401	151,131

in Westchester County parks. This finding alone implies that there is some risk of being bitten by *I. dammini* and of acquiring Lyme disease in seven of the eight parks sampled. Risk for Lyme disease is further implied by a previous study which detected spirochetes in 49.1 per cent of the nymphs and 53.6 per cent of the female *I. dammini* in Westchester County.[20]

The encounter distance used to quantify risk in this study is a more precise measure of risk for acquiring tick bites than is the total number of *I. dammini* collected in each park. Because the area of recreational parks may vary and sampling time may be restricted, it is difficult to achieve equal sampling effort at all sites. The encounter distance provides an index of tick abundance that allows for wide differences in total sampling effort.

The relatively higher encounter distances in the high use areas indicate that, for most parks, there is less risk of *I. dammini* bites in areas which are heavily used by the public. Fewer ticks in high use areas may be due to habitat alteration, such as buildings, roads, and maintained lawns which may discourage mammalian hosts or decrease tick survival. In studies of the lone star tick, *Amblyomma americanum*, areas with maintained vegetation have been found to have fewer ticks than wooded habitats.[21] Carey, *et al*,[22] found *I. dammini* more abundant in wooded than open areas. However, other studies have also determined that *I. dammini* may be prevalent in well maintained lawns.[23] Finding fewer ticks in high use areas could result from ticks attaching to humans and being removed from the host-seeking population.

The total public health risk from recreational activities in these parks may be significant (Table 4). For example, one park had both the highest patronage (151,131) and the highest *I. dammini* population index (146,401). With an *I. dammini* encounter distance of only 118 m, this park constitutes the greatest risk for tick exposure of all the parks studied. A similar situation existed at the park which had the lowest *I. dammini* encounter distance (36 m), and the third greatest park attendance (44,675).

In contrast to these extremes were the parks where the encounter distance were 3,333 m and infinity. Although the former had the second greatest number of patrons (84,606), it had a relatively low *I. dammini* population index (584). In the other, no *I. dammini* were found at all. Risk of exposure to *I. dammini* at these two parks is lowest of all those surveyed.

The presence of *I. dentatus* and *H. leporispalustris* in parks is not considered to be a public health threat. These species are not known to transmit Lyme disease or other illnesses to humans and they do not readily bite humans.[24] However, the presence of *I. cookei* may have public health significance as this species has been shown to be infected with Powassan encephalitis virus in the northeast[25] and is known to bite humans.[26,27]

The presence of *D. variabilis* in six of the eight parks sampled may also be significant because of this tick's role as the primary vector of Rocky Mountain spotted fever in New York State.[28] However, *D. variabilis* is much less common than *I. dammini* in this area,[24] and the incidence of Rocky Mountain spotted fever in this region is low.

Drag sampling is an efficient method for sampling *I. dammini* in large public recreational areas. It is a common tick collecting method[16,26,29,30] and has been applied in studies of Rocky Mountain spotted fever,[12,13,28,31] tularemia,[12,13] and Lyme disease.[20,23] To our knowledge, however, there is no previously published study describing a regimen suitable for quantitative sampling of nymphal *I. dammini* in large recreational areas.

Despite the efficiency of our sampling design, less than 1 per cent of each park was sampled for ticks. However, a mean of 21 hrs 40 mins spent at each park indicates that a surveillance project such as this requires considerable labor.

Unfortunately, practical methods to control *I. dammini* in areas the size of large recreational parks have not been developed. To reduce the incidence of Lyme disease in endemic areas, public health efforts should therefore be focused on identifying areas that present a high risk for acquiring tick bites. The public can then be educated to avoid tick infested areas and in ways to prevent ticks from biting. Educational efforts should include the posting of affected areas and the distribution of literature discussing the risk and prevention of tick bites and the symptoms of Lyme disease. We have demonstrated that public recreational parks should be an important target for this effort.

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Frequency of Worksite Health Promotion Activities

JONATHAN E. FIELDING, MD, MPH, AND PHILIP V. PISERCHIA, MS

Abstract: The first National Survey of Worksite Health Promotion Activities surveyed a random sample of all private sector worksites with 50 or more employees, stratified by number of employers, geographic location, and type of industry. The 1,358 completed interviews constituted a response rate of 83.1 per cent. Of responding worksites 65.5 per cent had one or more areas of health promotion activity with slightly more than 50 per cent of activities initiated within the previous five years. Overall prevalence by type of activity included health risk assessment (29.5 per cent), smoking cessation (35.6 per cent), blood pressure control and treatment (16.5 per cent), exercise/fitness (22.1 per cent), weight control (14.7 per cent), nutrition education (16.8 per cent), stress management (26.6 per cent), back problem prevention and care (28.5 per cent), and off-the-job accident prevention (19.8 per cent). Mean number of activities across all worksites was 2.1 and for worksites with activities, 3.2. Activity frequency increased with worksite size, was highest in the western region (2.34) and lowest in the northeast (1.96), and varied considerably by industry type. The majority of worksites paid the entire cost of these activities. (*Am J Public Health* 1989; 79: 16–20.)

Introduction

The worksite has been identified as an appropriate target for health promotion activities. Rationales for the workplace as an advantageous site for the efficient delivery of health promotion include: the large amount of time spent there by the majority of the population, the economic and other incentives for employers to invest in employee health promotion, the opportunity to mobilize peer pressure to help employees make desirable changes in health habits, as well as increasing reports of implementation of risk factor reduction at the worksite and some reports of program success.[1–11]

Response rates varied greatly in the three industry-specific[12–14] and six statewide surveys of health promotion activities performed between 1978 and 1986[15–20] and one survey (Colorado) included in their denominator only those worksites with an ongoing activity or planned programs. The reported percentage of worksites with at least one health promotion activity ranged from 17 per cent to 78 per cent. Surveys generally covered hypertension control, smoking cessation, weight control, nutrition, exercise/fitness, stress management and sometimes covered accident prevention, cancer risk reduction, mental health counseling, drug abuse education/referral, and cardiopulmonary resuscitation (CPR). In each survey, the prevalence of any activity and of most specific types of activities increased with worksite size. The results of one survey (California) supported the premise that employers are adopting these activities at an accelerating rate. This paper presents results of the first systematic nationwide assessment of worksite health promotion activities, including descriptive data on worksites surveyed, the employee population, and the frequency and nature of the activities.

Methods

The target population was all worksites in the private sector in the United States with 50 or more employees. The 1984 Dun and Bradstreet listing of private corporation worksites was used to identify a sample of worksites representative with respect to employee size, geographic location, and type of industry. The employee size reported by the worksite (which often differed from the size in the Dun and Bradstreet listing) was used for analysis and reporting purposes. Worksites identified as having 50 or more employees in the Dun and Bradstreet listing but reporting fewer than 50 employees were retained in the sample.

A sample of worksites was selected in each of two size categories. In the first category (50–99 employees), 600 worksites were selected with equal probability and allocated proportionally to the number of worksites in the universe in each of 24 classes defined by geographic region (four classes) and type of industry (six classes). In the second category (100 or more employees), 3,000 worksites were selected with equal probability and allocated proportionally to the number of worksites in the universe in each of 72 classes defined by geographic region (four classes), size of worksite (three classes), and type of industry (six classes). The sample selection used the Probability Minimum Replacement Method.[21]

Each sample site was randomly assigned to an interview list so that each list constituted a valid probability sample. Telephone interviews were conducted until a sufficient number of completed interviews was achieved to give coefficients of variation of about 5 per cent for the natural prevalence estimates. Once a list was begun, all the worksites on that list were contacted to ensure a valid probability sample. The 1,700 (Table 1) worksites in the final sample led to 1,358 completed interviews (320 for the employee size 50–99, and 1,038 for employee size ≥100).

Sampling weights were equal within each of the two size categories but, to avoid bias from nonresponse, the sampling weights of the respondents were adjusted so that their sum was equal to the number of worksites in the survey population. Because the sample of worksites was not a simple random sample and ratio adjustments were applied to the sample weights, special computer software was used for the computation of frequencies[22] and means.[23]

The instrument consisted of six sections. The introductory section confirmed that the appropriate worksite had been contacted, that the most appropriate person was being interviewed, and established the extent to which the employer, union, and employee group were providing financing and/or logistical support for these activities. The respondents at each worksite were queried about the demographic characteristics of the workforce, their own position and job respon-

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TABLE 1—Telephone Contact Results

	No. of Employees					
	<100		100+		Total	
	n	%	n	%	n	%
Original sample	400	100	1300	100	1700	100
Active at time of data collection*	400	100	1235*	100	1635	100
Completed interviews	320	80	1038	84.0	1358	83.1
Unable to complete interview†	67	16.8	190	15.4	257	15.7
Other‡	13	3.3	7	0.6	20	1.2

*65 worksites went out of business after sample selection but before name address files were received.
†Out of business, unable to contact, refusal.
‡Incorrect worksite size from Dun and Bradstreet list.

sibilities, and the presence of nine types of health promotion activities: health risk assessment, smoking cessation, blood pressure screening and treatment, exercise and physical fitness, weight control, nutrition education, stress management, back problem prevention and care, and off-the-job accident prevention. Worksites with at least one type of activity were asked to distinguish between passive activities, such as policy or information, and active employee participation, and to characterize the latter types, if present.

A lead letter to worksites introduced the survey and was followed by a computer-assisted telephone interview (CATI) with the person identified by that worksite as most familiar with its health promotion activities. Survey questions were displayed for the interviewer in program-controlled sequences on computer terminals, and responses were entered directly into the computer. The computer automatically performed checks for valid codes, consistency, and completeness.

There were 1,358 completed interviews, a response rate of 83.1 per cent (Table 1). The remaining 16.9 per cent included worksites which refused to participate, gave only partial data, could not be located, or were out of business. To complete an interview small worksites averaged three contacts and large worksites, four. Average interview time was 30 minutes.

Data analysis included frequencies and means, with frequency estimates of the universe obtained using RTIFREQS[22] and means using SESUDAAN,[23] both accessed through SAS software.[24]

Results

Worksites

Dun and Bradstreet data tended to overstate worksite size compared to worksite report. As a result, 137 of worksites reported having fewer than 50 employees despite contrary information in the sampling database. All surveyed worksites are included in the final sample to reflect the total universe based upon the original sampling frame. However, as worksites with fewer than 50 employees were not intentionally sampled, results for the 1–99 employee worksites cannot be generalized to the universe of worksites of that size.

Stratifying variables and frequencies based on worksite responses are displayed in Table 2. Although worksite size is an explanatory factor in the occurrence of all types of health promotion activities, unadjusted prevalences are used to accurately present baseline data by geographic region and industry type.

Workforce

Full-time employees (working 35 hours or more per week) made up over three-fourths of the workforce at 73 per cent (SE 1.4) of the worksites. Only 13.1 per cent (SE 1.1) of all worksites reported having no blue collar employees. The blue collar percentage of the population tended to decrease with increasing worksite size, with little regional variation.

Some union representation was reported by 25.5 per cent (SE 1.4) of worksites, with frequency varying directly with worksite size. A majority of represented employees was most frequent in the northeast United States (27.4 per cent SE 3.3). The percentage of blue collar workers, male workers, workers less than age 30, and workers represented by a union did not affect probability of a worksite having one or more health promotion activities.

Slightly more than one-half (51.6 per cent SE 2.3) of reported health promotion activities were initiated during the previous five years and 11.3 per cent (SE 1.6) within the prior 12 months. Longer duration activities included those in place 6–10 years (17.6 per cent SE 2.1), 11–20 years (10.4 per cent SE 1.3), and more than 20 years (8.6 per cent SE 1.4).

Health Promotion Activities

At least one type of health promotion activity was reported by 65.5 per cent (SE 1.7) of all worksites surveyed, with 2.1 (SE 0.1) activities per worksite for all responding worksites versus 3.3 (SE 0.1) for worksites with activities. For all worksites means increased from 1.37 (SE 0.1) in the smallest size (<100) to 4.64 (SE 0.3) in worksites with 750 or more employees. Interregional activity ranged from 1.96 (SE 0.2) (northeast region) to 2.34 (SE 0.2) (western region), considerably less than across industry type ranging from 1.47 (SE 0.2) in the "Other" category (Table 2) to 2.64 (SE 0.3) in the "Utilities Transportation Communications industry" (UTC).

The overall and worksite size specific prevalence rates for the nine activity types are shown in Table 2. For each, frequency increased with worksite size. Worksites with 750+ employees had an overall activity frequency of 87.6 per cent (SE 3.9), 1.6 times that of worksites with fewer than 100 employees, which was 55.5 per cent (SE 2.8). Frequency differences between the largest size (750+) and the next largest size (250–749) groups were generally greater than between the three smaller size groupings for each activity type. Frequency differences were small among regions but more pronounced among industries, and rank order by industry varied considerably by activity type.

Of all worksites, 29.5 per cent (SE 1.4) had some type of health risk assessment (HRA) activity. Health risk assessment activities were defined as any activity designed to measure employee health status or health risk, such as health risk appraisal, health examinations, physicals, or screenings. Periodic health or physical examination was available at 77.4 per cent (SE 2.3)* of these worksites and included blood pressure screening, 55.4 per cent (SE 3.1); tests of physical fitness, 15.0 per cent (SE 2.5); cancer screening, 19.5 per cent (SE 2.2); and blood tests for cholesterol, 28.3 per cent (SE 2.8) or sugar, 39.0 per cent (SE 2.8). A health risk assessment questionnaire was available at 24.0 per cent (SE 2.4) of all

*Frequencies for each specific activity are reported as a per cent of worksites with each type of activity.

TABLE 2—Frequency of Health Promotion Activities by Stratifying Variables [Prevalence in Per Cent and Standard Error (SE)]

Variables	% Total Worksite	Any Activity	Health Risk Assessment	Smoking Cessation	Blood Pressure Control	Exercise Fitness	Weight Control	Nutrition Education	Stress Mgmt.	Back Problem Prevention & Care	Off-the-Job Accident Prevention
Worksite Size											
<100 employees	50.2	55.5	18.5	30.1	8.8	14.7	8.6	9.3	15.1	19.2	12.8
		(2.8)	(2.0)	(2.4)	(1.7)	(1.8)	(1.7)	(1.7)	(1.7)	(2.0)	(1.6)
100–249	30.5	71.0	34.6	38.4	19.0	23.2	14.2	20.4	33.6	34.9	21.0
		(2.8)	(2.4)	(2.9)	(1.8)	(2.2)	(1.9)	(2.3)	(2.7)	(2.5)	(2.4)
250–749	12.9	80.4	41.1	40.0	24.0	32.4	22.5	21.9	37.6	40.7	33.1
		(2.7)	(2.9)	(3.4)	(2.7)	(3.0)	(2.5)	(3.0)	(3.2)	(3.3)	(3.2)
750+	6.6	87.6	66.1	57.6	48.8	53.4	47.9	47.7	59.6	46.4	36.8
		(3.9)	(5.4)	(5.4)	(5.8)	(5.6)	(5.8)	(4.9)	(5.3)	(5.3)	(5.8)
Region											
Northeast	23.6	62.9	26.8	36.0	15.6	23.0	11.0	16.3	23.5	23.7	20.5
		(3.8)	(3.0)	(3.4)	(2.6)	(1.6)	(1.7)	(2.9)	(2.5)	(2.7)	(2.9)
North Central	25.4	67.0	33.5	35.7	17.1	22.8	16.4	19.2	26.2	27.9	19.4
		(2.6)	(2.8)	(3.3)	(2.1)	(2.4)	(2.4)	(2.8)	(2.1)	(1.8)	(2.0)
South	32.6	61.8	25.7	30.1	16.4	19.3	14.1	16.1	24.7	31.7	21.4
		(3.3)	(2.0)	(2.4)	(1.6)	(1.8)	(1.7)	(1.9)	(1.7)	(2.6)	(2.3)
West	18.4	73.4	33.9	44.2	17.0	24.9	17.9	15.3	34.4	30.1	16.6
		(4.0)	(3.9)	(3.4)	(2.5)	(4.7)	(3.9)	(2.8)	(3.0)	(2.4)	(3.2)
Industry Type											
Manufacturing	29.4	65.1	30.1	34.3	18.2	16.7	11.0	10.9	19.8	30.9	20.0
		(3.2)	(2.9)	(3.0)	(2.0)	(2.0)	(1.7)	(1.7)	(2.0)	(2.4)	(2.4)
Wholesale/Retail	15.7	60.0	14.4	35.0	14.3	18.0	11.5	15.9	21.0	22.2	18.3
		(4.6)	(2.9)	(4.3)	(3.1)	(4.1)	(3.4)	(3.1)	(3.5)	(3.2)	(3.3)
Utilities/Transportation/Communications	3.6	77.9	42.6	21.2	17.5	19.6	13.7	17.0	45.1	35.6	51.3
		(6.3)	(9.6)	(5.6)	(6.0)	(6.7)	(7.2)	(6.6)	(5.7)	(6.7)	(7.8)
Financial/Real Estate/Insurance	6.2	60.1	31.4	32.9	13.5	31.3	20.3	19.7	39.1	11.9	10.7
		(6.7)	(5.9)	(7.8)	(4.0)	(5.8)	(4.5)	(6.1)	(6.3)	(3.5)	(3.5)
Service	37.5	70.7	35.3	42.0	17.9	28.5	20.1	23.6	33.3	30.3	16.7
		(2.4)	(2.3)	(2.2)	(1.7)	(2.3)	(2.2)	(2.4)	(2.2)	(2.0)	(1.8)
Other (includes Construction, Fishing, and Mining)	7.5	51.3	21.5	18.6	9.1	13.5	4.5	5.1	11.7	33.8	29.5
		(6.1)	(4.1)	(4.7)	(2.6)	(4.2)	(1.6)	(1.6)	(3.5)	(4.1)	(4.5)
All Worksites		65.5	29.5	35.6	16.5	22.1	14.7	16.8	26.6	28.5	19.8
		(1.7)	(1.7)	(1.5)	(1.1)	(1.2)	(1.2)	(1.3)	(1.2)	(1.2)	(1.3)

worksites with HRA activity, but at only 2.9 per cent (SE 0.15) of UTC worksites. None of the "Financial/Real Estate/Insurance" (FRI) or "Other" (OTH) worksites reported mammography screening.

Smoking cessation, the most prevalent activity present in 35.6 per cent (SE 1.5) of all worksites, included formal policies restricting smoking, 76.5 per cent (SE 2.4); information, 54.3 per cent (SE 2.8); classes or workshops, 17.5 per cent (SE 2.4); directly sponsored and self-help cessation materials, 49.7 per cent (SE 2.6). Formal policies restricting smoking were most prevalent at worksites with <100 employees, 84,6 per cent (SE 3.6) and in the "services" (serv) category, 81.8 per cent (SE 3.2). Participatory activities were individual counseling, 15.1 per cent (SE 2.0), group classes or workshops, 19.9 per cent (SE 1.8), and special events or innovations, 23.7 per cent (SE 2.4).

A 16.5 per cent (SE 1.1) prevalence was reported for high blood pressure control activities, ranging from 8.8 per cent (SE 1.7) at worksites with <100 employees, to 48.8 per cent (SE 5.8) at worksites of 750+ employees. Information was provided by 90.7 per cent (SE 2.1), treatment by 23.1 per cent (SE 2.8), referral to an outside physician for treatment by 63.3 per cent (SE 3.6), and special events or innovations by 26.2 per cent (SE 2.9).

Exercise and physical fitness activities were found at 22.1 per cent (SE 1.2) of all worksites surveyed, ranging from 14.7 per cent (SE 1.8) at smaller companies (<100) to 53.4 per cent (SE 5.6) at larger companies (750+). Of worksites with exercise fitness activities, 64.8 per cent (SE 3.0) provided information, 24.9 per cent (SE 2.54) individual counseling, 58.9 per cent (SE 3.5) sponsored group classes or workshops, 26.5 per cent (SE 2.9) subsidized health club membership, 25.4 per cent (SE 2.7) held special events, and 21.5 per cent (SE 2.6) had on-site exercise facilities or equipment for employee use. Worksites with 100–249 employees had the second highest frequency of providing subsidized membership to health clubs (29.9 per cent, SE 4.9) but the lowest frequency of exercise facilities or equipment available at the worksite (13.4 per cent, SE 2.9).

Weight control activities were reported by 14.7 per cent (SE 1.2) of worksites ranging from 8.6 per cent (SE 1.7) (<100 employees) to 47.9 per cent (SE 5.8) (750+ employees). Of these, information was provided by 76.8 per cent (SE 2.9) and self-help materials by 61.5 per cent (SE 3.9). Participatory activities included individual counseling by 42.9 per cent (SE 4.0), group classes or workshops, 51.4 per cent (SE 4.4), follow-up, 48.5 per cent (SE 4.8), and special events 23.0 per cent (SE 3.5). Individual counseling was most common in MAN at 48.2 per cent (SE 8.4) and "service" (Serv) categories at 50.7 per cent (SE 5.3), and group class/workshops

most frequent by far in FRI at 90.4 per cent (SE 6.3). Frequency of follow-up programs by industry type ranged from 50.8 per cent (SE 8.4) in MAN to 17.5 per cent (SE 13.2) in UTC.

The overall frequency for nutrition education activities was 16.8 per cent (SE 1.3) ranging from 9.3 per cent (SE 1.7) at the smaller firms (<100 employees) to 47.7 per cent (SE 4.9) at the larger firms (750+ employees), including information, 89.2 per cent (SE 2.6); group classes or workshops, 42.2 per cent (SE 3.7); individual counseling, 33.3 per cent (SE 3.1); and special events or innovations, 22.8 per cent (SE 2.9). Of worksites with a food service, healthy cafeteria food choices were provided at 51.3 per cent (SE 3.8); of worksites with vending machines, 33.7 per cent (SE 3.3) provided healthy choices.

Stress management activities were found at 26.6 per cent (SE 1.2) of all worksites, ranging from 15.1 per cent (SE 1.7) (<100 employees) to 59.6 per cent (SE 5.3) (750+ employees), with increasing frequency as worksite size increased. Activities included information, 80.7 per cent of worksites (SE 2.7); individual counseling, 39.3 per cent (SE 3.0); group classes or workshops, 58.1 per cent (SE 2.5); follow-up activities, 25.3 per cent (SE 3.5); and special events, 11.5 per cent (SE 1.6). Over 80 per cent reported organizational changes to reduce stress and two-thirds provided special places to relax. The north central region had the highest prevalence for most categories of stress management activity.

Activities to prevent and care for back problems were reported at 28.5 per cent (SE 1.2) of all worksites, ranging from 19.2 per cent (SE 2.0) (<100 employees) to 46.4 per cent (SE 5.3) (750+ employees), including information, 91.4 per cent (SE 1.7); group classes or workshops, 55.5 per cent (SE 3.0); and special events or innovations, 19.7 per cent (SE 2.1).

Almost 20 per cent (SE 1.3) of all worksites had off-the-job accident prevention activities, ranging from 12.8 per cent (SE 1.6) (<100 employees) to 36.8 per cent (SE 5.8) (750+ employees) directed at auto, home, and/or recreational injuries. Information was available at 75 per cent (SE 3.2) of the worksites, group classes or workshops at 27.8 per cent (SE 3.2), and special events at 24.1 per cent (SE 2.8). A formal seat belt policy was in effect at 61.5 per cent (SE 3.4) of worksites.

For nine types of activities surveyed, most companies in all size, region, and industry strata paid the full cost of all activities, including individual counseling at 69.9 per cent (SE 3.1), group counseling or workshops at 76.9 per cent (SE 1.9), and screening at 87.4 per cent (SE 1.9).

At 85.4 per cent (SE 1.9) of worksites with participation activities, all permanent employees were eligible, with a consistent pattern for all stratifying variables. The highest frequency of limiting eligibility to top management or executives was in FRI, at 15.8 per cent (SE 6.9), even though all permanent employees were eligible at 76.8 per cent (SE 7.6) of those worksites.

Coordination of health promotion activities was exclusively in-house at 41.8 per cent (SE 2.7), both inside and outside at 36.3 per cent (SE 2.1), mostly in-house at 14.1 per cent (SE 1.4), mostly outside at 1.7 per cent (SE 0.6), and exclusively outside at 5.4 per cent (SE 1.2). This pattern was consistent across all stratifying variables.

Only 15.9 per cent (SE 1.1) of all respondent companies indicated an identifiable budget allocation for health promotion activities; 55.3 per cent (SE 3.5) did not know the amount allocated for the current year. Distribution of known budget allocations were $1–999 at 7.6 per cent (SE 1.9); $1,000–9,999 at 14.2 per cent (SE 2.5); $10,000–99,000 at 17.3 per cent (SE 2.7); and $100,000–500,000 at 4.8 per cent (SE 2.0).

Discussion

Compared to results of statewide surveys performed during the five years prior to this survey, this national survey suggests large prevalence increases in overall health promotion activity, smoking cessation, nutrition education, and stress management; smaller increases were found in exercise/fitness; and no increase noted in weight control. Although methodological differences between surveys and lack of information on the life cycle of health promotion activities make any statements of trends tentative, the fact that greater than 50 per cent of the activities were in place for fewer than five years lends support to growing prevalence of these activities.

Employees at the 750+ worksites are at least twice as likely to have each type of activity as employees in <100 worksites, with prevalence ratios of about 5:1 for blood pressure control and treatment, weight control and nutrition education; 4:1 for stress management; and 3:1 for health risk assessment, exercise/fitness and off-the-job accident prevention. The increased likelihood of large worksites having one or more activities may be due to increased likelihood of dedicated personnel, benefits and health staff. Although regional differences are generally small, the south has a much lower prevalence for smoking cessation, which may relate to higher tolerance for smoking, concern for potential adverse regional economic effects of reduced tobacco consumption, and less state and local governmental action to restrict smoking. The south and northeast share a low relative prevalence of both health risk assessment and stress management. These differences do not correspond to known risk factor differences among employees at different size worksites.

No consistent rank order of prevalence exists among industry types although worksite size may be a confounding variable. A disproportionate number of UTC and FRI worksites are in the two largest categories, while the reverse is true for "wholesale/retail" and "other." Higher industry prevalence of some activities may reflect the type of work performed. MAN, OTH and UTC, associated with strenuous physical activity, have high rank order prevalence of back problem prevention and care activities. The need for many SER personnel to interact with customers may explain the high prevalence of smoking cessation activities, particularly smoking policies.

Prevalence rates for each of the two types of employee participation activities with the greatest evidence for effectiveness—smoking cessation and blood pressure control—are below 20 per cent, suggesting considerable expansion opportunities. Health risk assessment identifies reducible risks, and its results are frequently used to direct participants to specific behavioral change programs. Health risk assessment also supports program evaluation. Greater use of this instrument could therefore increase participation in other types of health promotion activities and permit better tracking of individual and group progress.

Although the contribution of worksite health promotion activities to achieving national health promotion objectives cannot be gauged from this study, all nine activity types support specific awareness and risk reduction objectives and their availability to tens of millions employees suggests that

the workplace be considered an important delivery site in any national risk education effort.

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The Paul B. Cornely Postdoctoral Program for Minority Scholars at U-MI SPH

The University of Michigan School of Public Health has completed its first round of admissions to the post-doctoral fellowship program for minority scholars, initiated last year. The first three postdoctoral fellows began their studies last September in this two-year program:

- Thomas LaVeist received his PhD in Sociology from U-MI in 1988, and his dissertation focused on Black-White differences in infant mortality.
- Dawn Smith, MD, recently completed her MPH in Public Health Policy and Administration from U-MI and is currently working on an AIDS intervention project.
- Dr. Verna Keith was an Assistant Professor of Sociology at the University of Texas, doing work on Black-White differences in physician utilization using the National Center for Health Statistics Health Interview Survey.

The program at U-MI School of Public Health has been named the Paul B. Cornely Postdoctoral Program for Minority Scholars. Dr. Cornely is an alumnus of the University of Michigan, having received his bachelor's degree there in 1928, his MD degree in 1931, and his DrPH in 1934. Dr. Cornely was the first Black American to earn a doctorate in a public health specialty. A long-time resident of Washington, DC, he was Director of the Howard University Hospital from 1947 to 1958; founder and first President of the District of Columbia Public Health Association; and the first Black President of the American Public Health Association.

Dr. Cornely has served on a number of White House Presidential Commissions and is an honorary fellow of the Royal Society of Health and of the American College of Hospital Administrators.

For further information about the post-doctoral fellowship, contact: Jan O'Connor, Woody Neighbors, or Marshall Becker, Dean's Office, University of Michigan School of Public Health, Ann Arbor, MI 48109.

National Trends in the Use of Preventive Health Care by Women

Diane M. Makuc, DrPH, Virginia M. Freid, BSN, MS, and Joel C. Kleinman, PhD

Abstract: This study investigates national changes between 1973–74 and 1985 in women's use of three preventive health services based on data from the National Health Interview Survey. Smoothed proportions of women with recent preventive care were estimated using weighted least squares for subgroups categorized by age (20–39, 40–59, 60–79), race (White, Black), income (poor, nonpoor), and year of interview. Older women and Black women experienced the largest increases in recent use of clinical breast examinations and Pap tests. Between 1973 and 1985 changes in recent breast examination ranged from zero for White women ages 20–39 years to a 23 percentage point increase (95% CI = 17,30) for Black women ages 60–79 years. A similar pattern was found for Pap testing. Changes between 1974 and 1985 in recent blood pressure testing ranged from zero for women ages 20–39 years to an 8 percentage point increase (95% CI = 6,10) for women ages 60–79 years. Despite the increases among older women, in 1985 recent use of breast exams and Pap tests remained lower among older women. Further, the poor remained less likely than the nonpoor to have recent preventive care (except blood pressure testing among older women). Most women without recent cancer screening tests had a recent physician contact, highlighting the need for greater emphasis on cancer prevention by health care providers. (*Am J Public Health* 1989; 79:21–26.)

Introduction

Since the late 1970s, there has been increasing interest in organizing health promotion and disease prevention activities for the American people. In a 1979 report, the US Surgeon General stressed that future reductions in morbidity and mortality were linked to the widespread adoption of preventive practices.[1] In 1979, the Canadian government also issued a landmark report highlighting prevention activities by recommending the replacement of the routine annual physical examination with a lifetime program of health promotion activities.[2] Canadian and US task forces are continuing to evaluate the scientific merit of specific preventive health measures to develop age- and sex-specific health screening recommendations.[3]

This article presents national changes between 1973–74 and 1985 in the use of three major preventive health services available to women: clinical breast examination, Pap smear testing, and blood pressure measurement. These widely accepted and commonly available preventive techniques have the potential to significantly decrease the levels of disease and disability among women. In this paper we investigate how changes over time in the use of these three preventive services vary by age, race, and income, and whether the subgroups at greatest risk of disease are as likely to receive preventive care as those at lower risk of disease.

Breast Cancer

Breast cancer, a leading cause of cancer deaths among women, accounted for 40,093 deaths in the United States in 1985.[4] Between 1973 and 1985, the age-adjusted breast cancer incidence rate has increased an average of 1 per cent each year, while the age-adjusted death rate for breast cancer has remained about the same.[5] The incidence of breast cancer among White women was almost 20 per cent greater than for Black women in 1985. However, Black women are at significant disadvantage to White women in terms of prognosis; only 62 per cent of Black women survive five years compared to 75 per cent of White women. Part of this difference is because Black women are less likely than White women to have localized disease at time of diagnosis (41 versus 49 per cent), but controlling for the stage of disease does not eliminate the disadvantage to Black women.[5]

Public health efforts have aimed to increase the detection of localized breast cancer through a three-pronged approach of breast self-examination, periodic clinical examination by a health care provider, and mammography.[6] The effectiveness of mammography and clinical breast exams in reducing mortality has been demonstrated among middle-aged and older women. However, the effectiveness of breast self-examination is less clear.[6] The American Cancer Society recommends a breast exam by a physician every three years between 20 and 40 years of age and an annual exam after age 40.[7] The American College of Obstetricians and Gynecologists (ACOG) recommends "regular periodic" breast exams.[8]

Cervical Cancer

In 1985, cervical cancer caused the deaths of 4,508 women in the United States.* Between 1973 and 1985 the age-adjusted incidence rate of invasive cervical cancer decreased 35 per cent, and the death rate for cervical cancer declined by 38 per cent.[5] Death rates have been declining at about the same rate, regardless of age or race.* In 1985 the age-adjusted incidence rate was twice as great for Black women as for White women and the age-adjusted death rate was two and half times as great.[5] Death rates and incidence rates for invasive cervical cancer also increase with age.

The appropriate frequency and ages for Pap test screening remain controversial.[9] Prior to 1980 the American Cancer Society actively advocated annual Pap tests for all women; in 1980 the Society changed its recommendation to an annual Pap test for two negative examinations and then at least every three years until the age of 65. A similar recommendation was made at a 1980 Consensus Development Conference at the National Institutes of Health. In contrast, the American College of Obstetricians and Gynecologists (ACOG) continues to recommend annual screening for "most women in the United States." The American Cancer Society is once again emphasizing annual Pap smears. The test frequency can be decreased at the discretion of the physician after three consecutive annual negative tests.[10] A national survey of

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*Unpublished data from the National Vital Statistics System, National Center for Health Statistics.

recently trained obstetricians and gynecologists found that over 97 per cent usually recommend annual Pap testing for women aged 20 to 59 years, and the majority of physicians (66–72 per cent) strongly disagree with the statement that "routine" Pap tests are not needed for women over the age of 65.[9]

Hypertension

About 22 million adult women ages 18–74 years (30 per cent) had elevated blood pressure readings (systolic blood pressure of at least 140 mmHg or diastolic blood pressure of at least 90 mmHg) during 1976–80.[11] The proportion of women with elevated blood pressure rises steadily with age, from 7 per cent of women 18–24 years to 64 per cent of women 65–74 years. Black women of all ages are at substantially higher risk of elevated blood pressure than their White counterparts.[11]

Mortality from heart disease and stroke has declined steadily since 1950 among women, partially due to decreases in undiagnosed and/or untreated hypertension.[11] Despite the gains of the past few decades, the detection of hypertension remains an important issue. The Joint National Committee on Detection, Evaluation and Treatment of High Blood Pressure has recommended that a blood pressure screening be conducted at every visit with a health professional or at least once every two years.[12]

Methods

This study uses data from the National Health Interview Survey (NHIS), a continuing household survey conducted by the National Center for Health Statistics (NCHS). The NHIS is a national probability sample of the civilian noninstitutionalized population of the United States. Each year the sample includes about 40,000 households. Information regarding health and other characteristics of household members is obtained through interviews in the household. Questions regarding the use of Pap tests and clinical breast examinations were included in the 1973 and 1985 surveys.[13,14] The use of blood pressure tests (taken by a physician or other health professional) was included in the 1974 and 1985 surveys.[14,15] In 1973 the entire NHIS sample received questions on use of preventive services (about 40,000 women). In 1974 and 1985, one-third of the NHIS sample received questions on preventive health care (about 15,000 women). In the 1973, 1974, and 1985 surveys interviews were completed for 96–97 per cent of all eligible households. Response rates to the three prevention questions vary from 96 to 98 per cent among those who received these questions.

Changes in use of preventive care are investigated for women in 12 subgroups based on age (20–39, 40–59, 60–79), race (White, Black), and income (poor, nonpoor). A poverty variable was constructed based on family income and family size using the 1973, 1974, and 1985 poverty threshold definitions.[16–18] Poor women were defined in this analysis as those with family income below 125 per cent of the poverty level for their family size. Information on family income was available for 93 per cent of the women in 1973 and 1974, and 89 per cent in 1985. Women with missing income data are excluded from estimates by income, but included in estimates which do not require data on income.

For each of the preventive services, trends in the per cent with a recent test are analyzed. Recent testing is defined as less than two years prior to the interview for Pap tests and breast exams and less than one year for blood pressure tests. In addition, changes in the per cent who have never had each test are examined. Smoothed proportions with a recent test were calculated for 24 subgroups of women based on age, race, income, and year of survey using linear models and weighted least squares.[19] The computer programs SURREGR[20] and SESUDAAN[21] were used to calculate observed proportions, variances, and covariances. These programs use Taylor series linearization to calculate variances that take into account the complex survey design of the NHIS. The 1985 NHIS sample weights were truncated at their 90th percentile because of the large variability in the sample weights for the 1985 NHIS health promotion and disease prevention supplement. The computer program GENCAT[22] was used to calculate smoothed proportions of women with a recent test by fitting linear models to the observed proportions. Because of the large differences in utilization according to age, model parameters were estimated separately by age using models nested within each age group. Saturated models (containing all possible interactions) were reduced by eliminating model parameters with a p value less than 0.10 in a hierarchical manner. Two additional interaction terms (with p values of 0.092 and 0.074, respectively) were eliminated because they had little effect on the smoothed per cents or the goodness of fit of the models. Results of goodness of fit tests and hypothesis tests for parameters in the final reduced models are found in the Appendix Table.

An additional analysis of the 1985 data partitions the per cent of women without a recent Pap test and breast exam into two groups, those who had a physician contact less than two years ago and those who did not. Contact with a physician includes contact with associated staff such as nurse practitioners and physician assistants. Apparent inconsistencies in response to questions regarding the time intervals since last physician contact and last Pap test and breast exam occurred for 1–1.5 per cent of women (e.g., a screening test was less than two years ago, but last contact with a physician or assistant was two or more years ago). The results presented in all tables and figures include these apparent inconsistencies. Analyses were repeated excluding inconsistent responses with virtually no difference in the results.

Results

Breast Exams

Between 1973 and 1985 the age-adjusted per cent with a recent breast exam increased from 66 to 70 per cent overall (Table 1). However, changes over time varied substantially by age and race with older women and Black women experiencing the greatest gains. To examine these trends in greater detail smoothed proportions of women with a recent breast exam were calculated according to age, race, income, and year (Figure 1). Among women 20–39 years of age, the per cent with a recent breast exam did not change for Whites, while it increased by 7 percentage points for Blacks (95% confidence interval = 4,11). Among middle-aged women (40–59 years) the per cent with a recent breast exam increased for women of both races, but to a greater extent for Black women than for White women. In addition, middle-aged poor women of both races showed somewhat greater gains than the nonpoor, so that increases ranged from only 2 percentage points among nonpoor Whites (95% CI = 0,4) to 15 percentage points among poor Blacks (95% CI = 8,22). The per cent with a recent breast exam increased substantially for all women 60–79 years of age, but the increases were greater for Black women (23 percentage points with 95% CI = 17,30)

TABLE 1—Per Cent of Women with Selected Preventive Services within the Past Two Years and Standard Errors according to Age and Race, 1973 and 1985

Age and Race	Breast Exam within the Past Two Years		Pap Test within the Past Two Years	
	1973	1985	1973	1985
Total[a]	65.5 (0.3)	69.6 (0.4)	63.8 (0.3)	64.8 (0.5)
Black[a]	61.7 (1.0)	74.8 (1.0)	59.9 (1.0)	70.5 (1.2)
White[a]	65.9 (0.3)	69.0 (0.5)	64.2 (0.3)	64.1 (0.5)
20–39 years	77.5 (0.4)	77.9 (0.5)	78.5 (0.4)	77.7 (0.5)
Black	77.0 (1.1)	83.9 (1.3)	78.0 (1.2)	84.6 (1.2)
White	77.6 (0.4)	77.0 (0.6)	78.6 (0.4)	76.7 (0.5)
40–59 years	62.1 (0.5)	66.0 (0.9)	59.5 (0.5)	59.9 (0.9)
Black	54.8 (1.4)	67.9 (2.2)	53.2 (1.6)	63.1 (2.4)
White	62.9 (0.5)	65.7 (0.9)	60.2 (0.5)	59.5 (0.9)
60–79 years	44.3 (0.6)	56.2 (0.9)	37.6 (0.6)	43.1 (0.9)
Black	39.1 (2.1)	64.5 (2.3)	31.2 (2.3)	50.4 (2.6)
White	44.7 (0.6)	55.4 (0.9)	38.2 (0.6)	42.4 (1.0)

[a]Age-adjusted by the direct method to the 1985 United States female civilian non-institutionalized population in 10-year age groups.

than for White women (10 percentage points with 95% CI = 8,12) (Figure 1).

In 1985, the per cent with a recent breast exam was 10–13 percentage points lower among the poor than the nonpoor for all age-race subgroups except one (among Black women 60–79 years this difference was only 5 percentage points with 95% CI = −2,13) (Figure 1). Despite the large increases in the use of breast exams among older women, they remain substantially less likely than middle-aged or younger women to have a recent exam in 1985 (56 versus 66 versus 78 per cent) (Table 1). Thus, in terms of age, frequency of breast examination is inversely related to risk of breast cancer. Black women were more likely than White women to have a recent breast exam in 1985 (75 versus 69 per cent), due to the greater increases in use of breast exams among Black women (Table 1).

All subgroups of women experienced large declines in the per cent who never had a breast exam (data available upon request to authors). Between 1973 and 1985, the age-adjusted per cent with no breast exam declined from 16 to 5 per cent. In 1985, the per cent who had never had a breast exam was higher among older women (11 per cent) than middle-aged or younger women (4 per cent). Older poor women were the most likely to have never had a breast exam (18 per cent).

Pap Tests

Trends in the use of Pap tests are similar to those for breast exams, particularly for younger women. This is not surprising because both procedures often occur at the same visit. In 1973 and 1985 the age-adjusted proportion with a recent Pap test was about the same (64 and 65 per cent, respectively) (Table 1). However, certain subgroups of women experienced substantial gains in the use of Pap tests over this period, particularly older women (60–79 years of age) and Black women.

Smoothed proportions of women with a recent Pap test according to age, race, income, and year (Figure 2) reveal trends similar to those for breast examinations. The slight decline in recent testing among young White women may reflect the 1980 changes in the American Cancer Society recommendations regarding the appropriate frequency of Pap tests.

As a result of the changes in Pap test use, in 1985 Black women were more likely than White women to have had a recent Pap test (71 versus 64 per cent), but within each race and age group the poor were still less likely than the nonpoor to have been tested recently (Table 1, Figure 2). In 1985, the per cent with a recent Pap test was 10–13 percentage points lower among the poor than the nonpoor for all age-sex subgroups except one (among Black women 60–79 years of age this difference was 4 percentage points, 95% CI = −3,11) (Figure 2). In addition, although the age differential was somewhat smaller in 1985 due to increases in Pap testing among older women, the per cent with a recent test remained substantially greater for younger women than for middle-aged or older women (78 versus 60 versus 43 per cent) (Table 1). Within each age group substantial variability in test use also remains. Poor White women were the least likely to have a recent Pap test and nonpoor Black women were the most likely (e.g., 47 per cent with 95% CI = 45,50 versus 66 per cent with 95% CI = 61,71 among middle-aged women) (Figure 2).

Although increases in recent Pap test use were limited to Black women and older White women, all subgroups of women have experienced large declines in the per cent who have never been tested (data available upon request to authors). Between 1973 and 1985, the age-adjusted per cent of women who have never been tested declined from 17 to 5

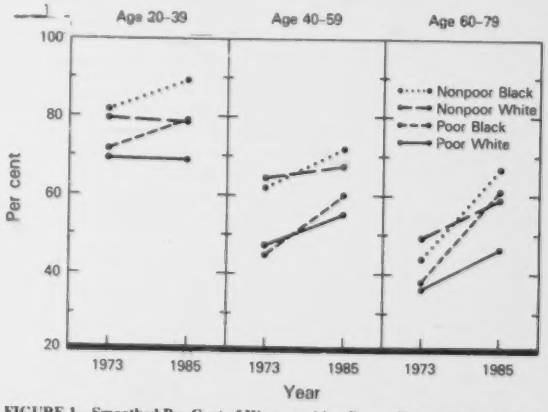

FIGURE 1—Smoothed Per Cent of Women with a Breast Exam within the Past Two Years, according to Age, Race, and Income, 1973 and 1985

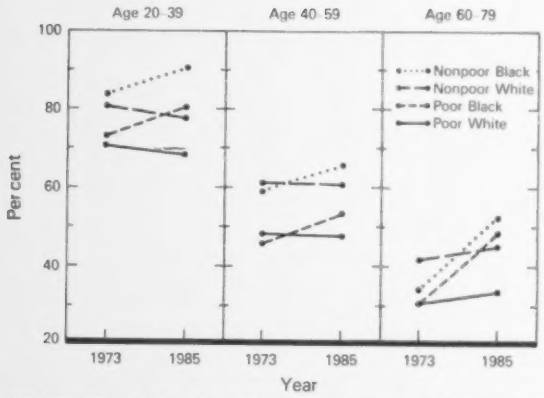

FIGURE 2—Smoothed Per Cent of Women with a Pap Test within the Past Two Years, according to Age, Race, and Income, 1973 and 1985

per cent. In 1985, the per cent who had never had a Pap test was highest among older poor women (18 per cent), the same as for breast exams.

Women without Recent Preventive Care

In Table 2, the per cent of women without a recent preventive screening test has been partitioned according to the time interval since the last physician contact. In 1985, the 30 per cent of women who did not have a recent breast exam included 22 per cent with a physician contact and only 9 per cent without a physician contact. Thus, among women without a breast exam during the past two years, 72 per cent (21.7/30.3) had a physician contact less than two years ago. The proportion of women without a recent breast exam who had a physician contact varied from 66 per cent for middle-aged women to 77 per cent for older women.

The results for Pap testing are similar to those for breast exams. The 35 per cent of women without a recent Pap test included 26 per cent with a physician contact and 9 per cent without a physician contact. Thus, among women without a recent Pap test, 75 per cent had a recent physician contact; this proportion ranged from 71 per cent among middle-aged women to 82 per cent among older women.

The results in Table 2 can also be used to partition the poverty differential in use of preventive care between women who have had a physician contact and those who have not. This provides some insight into whether the less frequent use of preventive care by the poor is primarily a problem of access to health care in general or a lack of emphasis on preventive care among women who have had some contact with the medical care system. For young women (20–39 years) and older women (60–79 years) most of the poverty differential in screening levels occurs among those who have had a recent physician contact. For example, among young women the 7 percentage point difference in use of breast exams between the poor and the nonpoor reflects a difference of 6 percentage points for women with a recent physician contact and only a 1 percentage point difference for women without a recent physician contact. Among middle-aged women the results are somewhat different. About half of the poverty differential in use of both Pap and breast exams occurs among women with a recent physician contact and the other half of the poverty differential is accounted for by the poverty differential among women without a recent physician contact.

Blood Pressure Measurement

Between 1974 and 1985, the age-adjusted proportion of women 20–79 years of age with a blood pressure measurement in the past year increased slightly from 76 to 79 per cent (Table 3). Examination of the smoothed proportions for the three age groups shows that the changes over time were the same for White and Black women and poor and nonpoor women. However, changes in the proportion with a recent blood pressure test increase with successive age groups. Older women (60–79 years) experienced the largest increase of 8 percentage points (95% CI = 6,10) (Figure 3). Unlike the results for Pap and breast examinations virtually all women had experienced a blood pressure test in their lifetime in 1974 and 1985.

In 1985, differences by race and income in the use of blood pressure testing were apparent for women under the age of 60. Consistent with their greater risk of hypertension, Black women were more likely than White women to have had a recent measurement. Among women under age 60, poor White women were least likely to be tested (72–73 per cent) and nonpoor Black women were most likely to be tested (83–86 per cent) (Figure 3). Among women 60–79 years of

TABLE 2—Per Cent of Women without Selected Preventive Services within the Past Two Years and Standard Errors according to Interval since Last Physician Contact, Age and Income, 1985

Age and Income	Per Cent without a Breast Exam within the Past Two Years			Per Cent without a Pap Test within the Past Two Years		
	Total[a]	and last physician contact ≥ two years ago	and last physician contact < two years ago	Total[a]	and last physician contact ≥ two years ago	and last physician contact < two years ago
Total	30.3 (0.5)	8.6 (0.3)	21.7 (0.4)	35.1 (0.5)	8.7 (0.3)	26.4 (0.4)
20–39 years	22.0 (0.5)	6.3 (0.3)	15.7 (0.5)	22.2 (0.5)	6.3 (0.3)	16.0 (0.5)
Poor	27.5 (1.2)	6.9 (0.6)	20.6 (1.0)	27.7 (1.2)	7.0 (0.7)	20.7 (1.1)
Nonpoor	20.4 (0.6)	6.1 (0.3)	14.3 (0.5)	20.9 (0.6)	6.0 (0.3)	14.9 (0.5)
40–59 years	33.9 (0.9)	11.4 (0.6)	22.5 (0.7)	40.1 (0.9)	11.6 (0.6)	28.5 (0.8)
Poor	43.8 (2.3)	16.5 (1.7)	27.3 (2.0)	49.3 (2.3)	16.5 (1.7)	32.7 (2.0)
Nonpoor	32.8 (1.0)	10.5 (0.6)	22.3 (0.8)	39.0 (1.0)	10.8 (0.6)	28.2 (0.9)
60–79 years	43.7 (0.9)	10.1 (0.6)	33.7 (0.8)	56.8 (0.9)	10.2 (0.6)	46.7 (0.9)
Poor	51.0 (1.6)	10.8 (1.1)	40.3 (1.5)	63.3 (1.6)	10.5 (1.1)	52.8 (1.8)
Nonpoor	40.6 (1.1)	9.3 (0.6)	31.4 (1.0)	54.8 (1.2)	9.5 (0.6)	45.3 (1.2)

[a]Excludes women (0.6 percent) with missing data on interval since last physician contact.

TABLE 3—Per Cent of Women with a Blood Pressure Measurement within the Past Year and Standard Errors according to Age and Race, 1974 and 1985

Age and Race	1974	1985
Total[a]	75.5 (0.4)	78.6 (0.4)
Black[a]	78.6 (1.3)	82.0 (0.8)
White[a]	75.0 (0.4)	78.2 (0.4)
20–39 years	76.5 (0.5)	77.2 (0.6)
Black	80.8 (1.4)	81.8 (1.2)
White	75.9 (0.6)	76.5 (0.6)
40–59 years	74.2 (0.6)	76.8 (0.7)
Black	79.0 (2.2)	80.0 (1.8)
White	73.6 (0.7)	76.4 (0.8)
60–79 years	75.6 (0.8)	84.2 (0.7)
Black	74.8 (3.0)	85.1 (1.7)
White	75.7 (0.8)	84.1 (0.7)

[a]Age-adjusted by the direct method to the 1985 United States female civilian noninstitutionalized population in 10-year age groups.

age, 84 per cent reported a recent blood pressure test in 1985, regardless of race or income (Figure 3). In contrast to the findings for Pap and breast examinations, the proportion with a recent blood pressure test did not decrease with advancing age. In 1985, young (20–39 years) and middle-aged women (40–59 years) were equally likely to be tested, and older women were the most likely to have had their blood pressure tested recently.

Discussion

Between 1973–74 and 1985, some encouraging changes occurred in the use of preventive health screening. The proportion never screened for breast or cervical cancer declined sharply for all subgroups of women. Recent screening tests for cervical cancer, breast cancer, and hypertension increased among older women who are at increased risk of these diseases. In addition, recent screening tests for cervical and breast cancer increased for Black women who are at higher risk for cervical cancer and are more likely to have advanced stages of breast cancer at the time of diagnosis.

Increases in preventive care for Black women occurred in all age groups and for both poor and nonpoor Black women. The consistency of the increases in preventive care across all subgroups of Black women, in contrast to the much

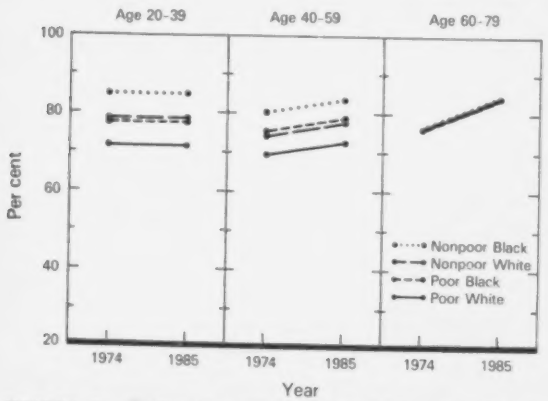

FIGURE 3—Smoothed Per Cent of Women with a Blood Pressure Test within the Past Year, according to Age, Race, and Income, 1974 and 1985

smaller changes (or no change) observed for the corresponding subgroups of White women, is noteworthy. The greater use of preventive screening tests among Black women in 1985 contrasts sharply with racial differences in the use of early prenatal care. Black mothers are less likely than White mothers to begin care during the first trimester of pregnancy, regardless of age or educational attainment.[23] Racial differences for other measures of ambulatory care use were minimal in 1985.**

Increasing emphasis on the importance of health promotion and preventive practices between 1973 and 1985 may have contributed to the increases in use of preventive care. Several large screening programs were established including the breast cancer detection demonstration project,[24] the community hypertension evaluation clinic program,[25] and the national high blood pressure education program.[26] Such programs can heighten awareness of the need for prevention among both health care providers and consumers. Also during this period, women's magazines and the popular press have publicized the importance of preventive health screenings.

Despite the gains since 1973 among Black and older women, inequities remain in the use of preventive care. In 1985, poor women of all ages and races were less likely than their nonpoor counterparts to have recent preventive screenings (except blood pressure measurement among older women). Older women remain less likely to have had a recent breast exam or Pap test, despite the fact that the incidence of these diseases rises sharply with age. In addition, poor women and older women were more likely to never have received breast or Pap screening in their lifetime. Lower levels of recent Pap testing with increasing age may be partially related to the high prevalence of hysterectomies. Data from the 1982–84 NHANES I Epidemiologic Follow-up Study show that 30 per cent of women over the age of 40 reported having had a hysterectomy.*** However, the American College of Obstetricians and Gynecologists recommends continued cytologic testing even for women who have had hysterectomies.

Most women who did not have a recent Pap test or breast exam did have a recent contact with a physician. Further, the lower use of screening tests among poor women than nonpoor women is accounted for in large part by women with a recent physician contact. However, among middle-aged women a substantial part of the poverty differential in screening is also accounted for by women without a recent physician contact. The relatively high proportion of middle-aged poor women without a recent physician contact (17 per cent) may reflect a lack of health programs available to poor women in this age group.

Women with recent physician contacts may not receive preventive care for a variety of reasons. Not every contact with a physician presents the opportunity for cancer screening tests. Some physician contacts may be visits to specialists, e.g., ophthalmologists, who do not provide this type of service. Other physician contacts which may be unlikely to include preventive screening tests include visits for acute illnesses, particularly emergency room visits.

**Unpublished data from the National Health Interview Survey, National Center for Health Statistics, and the National Ambulatory Medical Care Survey, NCHS.

***Unpublished data from the National Health and Nutrition Examination Survey (NHANES I) Epidemiologic Follow-up Study, National Center for Health Statistics.

Further efforts to encourage use of preventive care are warranted, especially for selected subgroups of women. In particular, in 1985, the majority of older White women, half of older Black women, and half of middle-aged poor women did not have a recent Pap test. Perhaps even more importantly, among poor White women, the majority of those 60–79 years of age and almost half of those 40–59 years of age did not have a recent breast exam. These low levels of screening for breast and cervical cancer contrast sharply with the much higher levels of blood pressure testing among these subgroups of women, and are due in large part to a lack of emphasis on cancer prevention for women who have had contact with the health care system.

APPENDIX TABLE
Final Models Used to Calculate Smoothed Per Cents with Recent Preventive Care according to Age, Race, Income and Year

Model and Hypothesis Test		X^2	df	p-value
Breast exam within past two years				
Goodness of fit		4.12	7	.766
Age 20–39:	RY = 0	14.94	1	<.001
	I = 0	153.82	1	<.001
Age 40–59:	RY = 0	7.05	1	.008
	IY = 0	3.96	1	.047
Age 60–79:	RY = 0	14.65	1	<.001
	RI = 0	4.20	1	.041
Pap test within past two years				
Goodness of fit		6.80	8	.559
Age 20–39:	RY = 0	24.66	1	<.001
	I = 0	171.00	1	<.001
Age 40–59:	RY = 0	7.46	1	.006
	I = 0	135.62	1	<.001
Age 60–79:	RY = 0	20.69	1	<.001
	RI = 0	4.75	1	.029
Blood pressure test within past year				
Goodness of fit		11.68	15	.703
Age 20–39:	R = 0	53.51	1	<.001
	I = 0	32.93	1	<.001
Age 40–59:	R = 0	16.67	1	<.001
	I = 0	10.63	1	.001
	Y = 0	8.42	1	.004
Age 60–79:	Y = 0	61.16	1	<.001

NOTE: Models are hierarchical and include main effects for all interactions shown. Tests for main effects included in the models are shown when interactions are not present. R = race; Y = year; I = income.

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Resistance to Sexual Assault: Who Resists and What Happens?

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Abstract: To determine who resists sexual assault and what happens, data were examined from a probability sample of 3,132 adult community residents of Los Angeles, California. Seventy-five per cent of the respondents reporting an assault (n = 365) indicated that they had attempted to resist their most recent assault; talking was the most frequently used resistance strategy. The strongest predictor to emerge in the multivariate analyses of resistance was timing of assault: respondents assaulted only in childhood were less likely to resist than either respondents assaulted only in adulthood, or respondents assaulted in both phases. Univariate analyses indicated that resistance reduced the probability of sexual contact, however multivariate analyses suggested that assailant use of force was the most important determinant of assault outcome. (*Am J Public Health* 1989; 79:27–31.)

Introduction

Legal thought has been concerned with the topic of resistance to sexual assault because victim resistance has been considered the proof that sexual overtures are unwanted. Victims of sexual assault are unique in this respect—victims of no other violent crime are expected to meet their assailant with resistance.[1] While most states no longer require proof of resistance in criminal rape proceedings, societal norms still emphasize resistance as one way to legitimize a victim's claim that sexual contact was unwanted.

Research on resistance to sexual assault has focused largely on the relative effectiveness of various strategies in deterring assault and physical injury. Conclusive statements about strategy effectiveness have been preempted, however, by the variability in definitions of resistance across studies. For example, screaming has been alternatively grouped with other verbal strategies or with more aggressive physical strategies. Studies of resistance have been limited also by their reliance on respondents who report assault to the police or who utilize rape crisis centers. Biased estimates of the probability of resistance may emerge because reported assaults are known to differ in important ways from non-reported assaults,[2] and the dimensions on which these assaults differ may be related in turn to the use of specific resistance strategies. Successful resistance strategies (those deterring assault) may be under reported as well.[3–5] Despite these biases and definitional problems, the empirical research converges to support one important conclusion: resistance reduces the probability of an assault being completed.

Fewer studies have examined the relation between resistance and physical injury other than rape. Two investigations suggest that resistance may be costly to the victim in terms of physical injury.[6,7] In contrast, other researchers[8,9] conclude that resistance does little to increase one's chances of injury. One analysis that took timing into account addresses this apparent contradiction.[10] While a variety of resistance strategies are associated with increased victim injury, the chronology shows that victim resistance increases in response to injury. These and other data[11] led Furby and Fischhoff[3] to suggest that more assertive strategies are associated with greater chances of injury, but these strategies are consequences, not causes, of increased injury.

The research reported here differs from most studies of resistance because the sample is community-based and the respondents are not selected specifically because of their sexual assault experience. Also, data are gathered from men, as well as women, and from Hispanics, as well as Whites who are not Hispanic. (Prevalence of assault for this sample is reported elsewhere.[12,13]) The first aim of this research is to determine the degree to which demographic characteristics and circumstances of assault are associated with resistance strategy. One prior study found that female rape victims reported greater resistance if they knew their assailant, if no weapon was present, and if they displayed certain personality characteristics (such as dominance and internal locus of control).[14]

Like other investigations, this research also examines the relation of resistance to assault outcome, including physical injury. Multivariate analyses allow for a test of independence among the variables associated with resistance or outcome.

Methods

The data are from the Los Angeles Epidemiologic Catchment Area (ECA) study, one of five sites of the National Institute of Mental Health (NIMH)-initiated collaborative ECA program. Respondents at the Los Angeles site were also asked a series of questions regarding lifetime sexual assault. The design and methodology of the ECA surveys are described elsewhere.[15,16]

Sample

The ECA sample consists of 3,132 adults (18 years of age and older) from two catchment areas. One catchment area contains predominantly Hispanic Americans (83 per cent) and the other contains largely non-Hispanic Whites (63 per cent) but also many Hispanics (21 per cent). Eighty-seven per cent of the Hispanic Americans in the two catchment areas are of Mexican origin. Respondents chose whether to take the interview in English or Spanish. The survey instrument allowed easy switching from one language to another.

The sample, stratified by catchment area, was selected using a two-stage probability technique. One adult from each household was randomly selected for inclusion using the Kish procedure.[17] Respondents were interviewed in person, with an overall completion rate of 68 per cent. Among those consenting to the interview, less than 1 per cent (21 of 3,132 respondents) refused to answer the filter question on experience with sexual assault.

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The data presented here represent the experiences of the 365 respondents in the Los Angeles ECA sample who provided data about resistance in response to their lifetime experience with pressure or force to have sexual contact. The demographic characteristics of these respondents are presented in Table 1.

Instrument

Lifetime sexual assault was assessed using the question: "In your lifetime, has anyone ever tried to pressure or force you to have sexual contact? By sexual contact I mean their touching your sexual parts, your touching their sexual parts, or sexual intercourse?" Those answering affirmatively were asked if they had ever been forced or pressured for sexual contact before age 16 (childhood sexual assault). Detailed information was collected from all assaulted respondents concerning the most recent assault. Among other questions, respondents were asked: time elapsed since assault; the relation of the assailant to the respondent; the kind of pressure or force used; whether physical harm occurred and when; the type of sexual activity that occurred; and whether resistance strategies were used. (A more detailed description of this instrument is reported elsewhere.[12])

Operationalization of Variables

The data presented here focus on two aspects of resistance to assault: the correlates of resistance and the relation of assault resistance with reported outcome. Respondents reporting any sexual assault were asked the following question: "We are interested in the kinds of situations in which people may decide to resist sexual contact. The most recent time you were pressured or forced into sexual contact did you try to resist?" Respondents who indicated they had tried to resist were then asked: "What did you do?" Interviewers recorded the respondents' answers verbatim.

Correlates of resistance examined in this research include respondent demographic characteristics and assault circumstances:

- The *demographic variables* are: gender, age at interview (dichotomized at younger than 40 years and 40 years or older), ethnicity (Hispanic or other White), and education (dichotomized at the median years of education within ethnic group, i.e., 10 among Hispanics, 14 among other Whites).
- *Circumstances surrounding the assault* are: whether the respondent had experienced one compared to two or more lifetime assaults; whether the most recent assault had occurred during childhood (below age 16) or during adulthood (this group was subdivided into respondents assaulted only in adulthood and those assaulted in adulthood with at least one prior assault in childhood); the time elapsed since this assault; what kind of pressure or force the assailant used (described below); and the respondent's relation with the assailant (described below).

The question on pressure or force asked the respondent to indicate from a list all strategies that applied. For purposes of analyses, these assailant strategies were classified as either *pressure* (they tried to talk you into it; you were afraid they wouldn't like or love you; they bribed you) or *force* (they said they would hurt you; they scared you because they were bigger or stronger; they pushed, hit, or physically restrained you; they threatened you with a weapon; they drugged you or got you drunk; they physically harmed or injured you). Those indicating physical harm were also asked the following three questions: "Were you hurt *before* sexual activity took place?"; "Were you hurt *during* sexual activity?"; "Were you hurt *after* sexual activity had taken place?"

The respondent's relation to the assailant was classified for analyses as stranger, acquaintance (includes friend), intimate (lover, spouse), relative, and other. Regarding assault outcome, respondents were asked: "What kinds of sexual activity resulted?" A list of applicable responses included: none, attempted sexual contact only; they touched your sexual parts (genitals or breasts); they had you touch their sexual parts; intercourse; something else.

TABLE 1—Demographic Characteristics of Respondents Reporting Sexual Assault, Los Angeles Epidemiologic Catchment Area Site

Sex, Age	Mexican American	Other White	Other	Total
Men[a]				
Age (years)				
18–39	28	51	11	90
40+	6	24	3	33
Total	34	75	14	123
Women[b]				
Age (years)				
18–39	46	110	11	167
40+	16	52	2	70
Total	62	162	13	237
Total	96	237	27	360

[a]The data for two men are omitted because they did not report their age.
[b]The data for three women are omitted because they did not report their age.

Results

Resistance Strategies

All responses to the open-ended question on resistance strategies used in the most recent assault were reviewed and content analyzed in order to develop coding categories. Up to two resistance strategies per respondent were coded. Very few respondents reported more than two resistance strategies or refused to answer this question (about 1 percent for each). This content analysis yielded six categories: talking; reasoning; saying no sexual interest; loud or angry talking; fleeing; and physically fighting.

Other resistance strategies were infrequently occurring responses (less than 4 per cent) that could not be classified in these six categories, such as "throwing up" or "making him take me home." Some respondents were classified into the no resistance category, even though they reported they had attempted to resist assault, if their resistance strategy indicated no apparent resistance. Three raters (JMG, JAS, JMS) independently classified each of the open-ended responses. All three raters agreed on the category 90 per cent of the time; two of three raters agreed 98 per cent of the time. The raters discussed the responses for which there was complete disagreement and consensus was achieved.

Table 2 presents the frequencies for reported use of each of the six categories of resistance strategies. These data are presented for all respondents, for men, and for women. As shown in Table 2, talking was the most frequently used resistance strategy (27 per cent), followed by physically fighting (23 per cent). Saying no sexual interest or loud/angry talking were used less often (3 and 4 per cent, respectively), while fleeing (15 per cent) and reasoning (11 per cent) were used more often. Combining all resistance strategies, 75 per cent of the respondents indicated they had attempted to resist their most recent assault.

TABLE 2—Frequency[a] of Use of Resistance Strategies for All Assaulted Respondents, for Women, and for Men

Resistance Strategy	All Respondents n = 365 Per Cent[b]	(n)	Women n = 240 Per Cent[b]	(n)	Men n = 125 Per Cent[b]	(n)
Talking	27	(91)	23	(49)	35	(42)
Reasoning	11	(36)	10	(21)	13	(15)
Saying no sexual interest	3	(11)	3	(6)	4	(5)
Loud/angry talking	4	(15)	7	(15)	0	(0)
Fleeing	15	(49)	17	(36)	11	(13)
Physically fighting	23	(77)	30	(64)	11	(13)
No resistance	25	(84)	22	(47)	31	(38)

[a] Weighted for sample design and non-response.
[b] Percentages do not sum to 100 because respondents could identify multiple strategies.

Correlates of Resistance

To examine the correlates of resistance, the strategies presented in Table 2 were aggregated into verbal resistance (talking, reasoning, saying no sexual interest, loud/angry talking), physical resistance (fleeing, physically fighting), and no resistance. Respondents who indicated use of both verbal and physical resistance were classified with the physical resisters. The findings were unchanged if this latter combination strategy was considered separately.

The rates (per cent reporting) of each resistance strategy were first examined as a function of gender, ethnicity, age, and educational attainment. These rates were calculated using SESUDAAN, a specialized SAS subprogram which calculates rates and asymptotic approximations of exact standard errors in clustered samples.[18] Analyses of ethnicity and educational attainment included only the Hispanic and other White respondents, as only these groups were sufficiently represented in the sample to permit comparison of rates and computation of median educational level.

Age at interview, ethnicity, and educational attainment were unrelated to the use of resistance strategy. Women were more likely to use physical (or combination) resistance strategies than men (46.6 per cent, 99 per cent confidence intervals [CI] = 36.9, 56.3 versus 22.0 per cent, 99 per cent CI = 11.8, 32.2). In contrast, men were more likely than women to use verbal strategies (46.9 per cent, 99 per cent CI = 35.1, 58.8 versus 31.1 per cent, 99 per cent CI = 22.3, 40.7) or no strategies (31.1 per cent, 95 per cent CI = 22.0, 40.2 versus 21.8 per cent, 95 per cent CI = 17.0, 26.7).

Turning to the circumstances surrounding the most recent sexual assault, the number of lifetime assaults, the time elapsed since the most recent assault, and the respondent's relation to the assailant were all unrelated to resistance strategy. Respondents whose most recent assault was during childhood (at age 15 or younger) were more likely to offer no resistance (46.9 per cent, 99 per cent CI=31.1, 62.8) than either respondents assaulted only in adulthood (21.6 per cent, 99 per cent CI = 14.1, 29.2) or those assaulted in both adulthood and childhood (16.5 per cent, 99 per cent CI = 4.7, 28.4). Additionally, respondents assaulted only in childhood were less likely to use verbal resistance (17.7 per cent, 99 per cent CI = 4.4, 31.0) than the two adult groups (44.4 per cent, 99 per cent CI = 33.9, 54.8 and 32.2 per cent, 99 per cent CI = 19.2, 45.2, respectively). Use of physical or combination of strategies was equivalent for those assaulted only in childhood (35.4 per cent) or only in adulthood (34.0 per cent), and was more prevalent among respondents assaulted during both life phases (51.2 per cent, 99 per cent CI do not overlap with other groups). Respondents whose assailant used pressure were more likely to use verbal strategies (46.0 per cent, 99 per cent CI = 35.8, 56.3) than were respondents whose assailants used force (21.3 per cent, 99 per cent CI = 9.7, 32.9). Complementary findings emerged for physical or combination strategies: these strategies were more common among those experiencing force (56.9 per cent, 99 per cent CI = 40.8, 73.0) than pressure (26.3 per cent, 99 per cent CI = 15.7, 36.8). The probability of offering no resistance was unrelated to the assailant's use of pressure (27.7 per cent) or force (21.8 per cent).

To control for the possibility of shared variance among the correlates of resistance, all variables were examined simultaneously in a logistic regression analysis, using the SAS CATMOD procedure.[19] For this analysis, resistance was categorized as a dichotomous (yes,no) variable. As shown in Table 3, developmental timing of assault was the strongest predictor of resistance. Respondents assaulted only in childhood were less likely to resist than either respondents assaulted only in adulthood (odds ratio [OR] = .18, 95 per cent CI = .60, .52) or respondents assaulted in both life phases. Inclusion of all two-way interaction terms in the model did not increase the ability to predict resistance.

Relation of Resistance to Assault Outcome

Assault outcome was examined in three ways. First, outcome was dichotomized as aggregated categories of: contact (they touched your sexual parts, they had you touch their sexual parts, intercourse, something else involving contact, e.g., oral sex) compared to no contact (category of none, attempted sexual contact only). Second, outcome was dichotomized as intercourse (oral, anal, or vaginal) compared to all other outcomes, and third, outcome was dichotomized into categories of physical harm compared to no harm. Data on physical harm were obtained from the question on the type of pressure or force utilized by the assailant; one response option was, "They physically harmed or injured you."

The rates (per cent reporting) of each resistance strategy were examined as a function of contact, intercourse, and physical harm. Those experiencing contact were more likely to employ no resistance strategy (31.2 per cent, 99 per cent CI = 23.5, 38.9) than those with no contact (13.1 per cent, 99 per cent CI = 5.4, 20.9). The two groups were equally likely to employ physical plus combination strategies (38.8 per cent compared to 37.7 per cent), and those experiencing no contact reported a greater use of verbal strategies than those with contact (48.1 per cent, 95 per cent CI = 36.6, 59.7 versus 31.1 per cent, 95 per cent CI = 25.2, 37.1). Data on intercourse as an outcome compared to the aggregated categories of all other sexual activity showed no relation with resistance.

TABLE 3—Predictors of Resistance to Sexual Assault: Multiple Logistic Regression Analysis

Variables	Coefficient	S.E.	Odds Ratio	95% Confidence Interval Limits	
				Lower	Upper
Sex (Female, Male)	−0.09	0.19	0.92	0.63	1.35
Age at Interview (40 or Above, Below 40)	−0.40	0.21	0.67	0.45	1.01
Ethnicity (Other, White, Hispanic)	0.14	0.19	1.14	0.79	1.64
Education (Above Ethnic-Specific Median, Below)	−0.12	0.18	0.89	0.62	1.27
Number of Lifetime Assaults (2 or more, 1)	−0.28	0.22	0.76	0.50	1.16
Timing of Assault					
Childhood only vs Adulthood only	−1.69	0.54	0.18	0.06	0.52
Childhood only vs. Child and Adulthood	−2.15	0.70	0.12	0.03	0.45
Assault Recency (Above Median, Below)	−0.10	0.21	0.90	0.60	1.35
Assailant Method (Force, Pressure)	0.11	0.18	1.12	0.78	1.61
Relation to Assailant (Known, Stranger)	0.01	0.23	1.01	0.65	1.57

Likelihood ratio = 0.67

To determine whether resistance was related to assault outcome independently of other potential correlates of outcome, resistance, demographic variables, and characteristics of the assault were all included as predictors of assault outcome (contact, intercourse) in logistic regression analyses. The most important determinant of contact was assailant use of force, with force (OR = 1.62, 95 per cent CI = 1.21, 2.17) more likely to result in contact than pressure. The probability level for the likelihood ratio chi-square for this regression model was very low (.01), suggesting that the model was not a good fit. Inclusion of all two-way interaction terms did not improve model fit. Thus, these results should be viewed as tentative.

Predicting intercourse with these variables resulted in the same effect for assailant use of force. Specifically, force was (OR = 1.41, 95 per cent CI = 1.06, 1.87) more likely to result in intercourse than pressure. The other independent predictor was time of assault: those assaulted only in adulthood (OR = 1.83, 95 per cent CI = 0.05, 0.50) and those assaulted both in adulthood and childhood were (OR = 1.70, 95 per cent CI = 0.05, 0.62) more likely to have intercourse than those assaulted only in childhood. Similar to the regression predicting contact, the probability level of the likelihood ratio chi-square was quite low.

Turning to the data on physical harm, respondents who reported that they were physically harmed were more likely to report the use of physical plus combination resistance strategies than respondents reporting no harm (67.6 per cent, 99 per cent CI = 43.5, 91.8 compared to 35.3 per cent, 99 per cent CI = 26.9, 43.6), and they were less likely to report verbal resistance than respondents who were not harmed (16.1 per cent, 99 per cent CI = 0.2, 32.2 compared to 39.2 per cent, 99 per cent CI = 31.0, 47.3). A similar trend emerged for no resistance strategy (16.2 per cent, 95 per cent CI = 2.7, 29.8 compared to 25.6 per cent, 95 per cent CI = 20.2, 30.9). To further explore the relation between physical harm and sexual assault, data concerning when the physical harm occurred were examined. Of the 34 respondents who reported physical harm, 29 answered the questions about when harm occurred. For 23 respondents (79 per cent), harm took place before sexual activity; 6 respondents (21 per cent) reported being hurt during or after sexual activity exclusively. Data were not gathered on when resistance was offered.

Discussion

These data, gathered from a probability sample of community respondents, are the first to address comprehensively, and with multivariate analyses, the question of who resists sexual assault. In both univariate and multivariate analysis, individuals whose most recent assault occurred in adulthood (age 16 or older) are more likely to report they resisted than individuals whose most recent assault occurred in childhood. Several factors may contribute to the lower likelihood of resistance among children and adolescents. Children may be more trusting of others and less able to distinguish at onset an act of affection from an attempt at sexual contact. This is exacerbated by less restrictive societal norms for touching a child (in a non-sexual way) compared to touching an adult. Further, children are socialized to be deferent and obedient, particularly to those in positions of authority or power. Respondents assaulted only in adulthood were equally likely to resist as those assaulted in adulthood and childhood, but the strategies of the two groups differed. Respondents assaulted during both life phases reported greater use of physical strategies and less use of verbal strategies.

Univariate analyses show that resistance strategy matches assailant strategy, i.e., use of force by the assailant is associated with physical resistance by the victim, and use of pressure by the assailant is associated with victim verbal resistance. The finding that women are more likely than men to utilize physical resistance probably reflects that for those describing an adult assault, harm or threat of harm is more frequently reported by female than male respondents.[13]

Regarding the relation between resistance and outcome of assault, our univariate analyses corroborate prior research in

showing that resistance, particularly verbal, reduces the probability of sexual contact. Physical resistance, on the other hand, is associated with increased likelihood of contact. These effects did not emerge in multivariate analyses, suggesting that the relation of resistance to outcome is a function of other variables measured in this study, most likely assailant use of force. Also relevant to assault outcome is that physical harm relative to no harm was associated with greater use of physical resistance strategies. The majority (79 per cent) of those harmed indicate that harm commenced prior to sexual activity; thus like other investigations,[10,11] these findings imply that physical resistance is more likely to be a consequence of the sexual assault rather than provoking further injury. Taken together, these results suggest that assailant use of force is more effective than verbal pressure in accomplishing the assailant's aim, *and* that use of force acts as a stimulus for victim physical resistance.

It is important to note, however, that these data do not describe the resistance experiences of persons who were physically harmed to the point of incapacitation or death.

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"Children at Risk" Conference Set for April in Switzerland

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Breastfeeding Trends in the Philippines, 1973 and 1983

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Abstract: This paper examines comparable national surveys of breastfeeding from the Philippines carried out in 1973 and 1983. The probability of breastfeeding at selected infant ages is estimated, using the weighted life table. The conclusions are that a 5 per cent decline in the proportion of infants ever breast-fed occurred during the referenced period, and that median length of breastfeeding remained essentially the same. (*Am J Public Health* 1989; 79:32–35.)

Introduction

In recent years, reports of widespread declines in breastfeeding in developing countries have generated much interest, especially among policy decision makers.[1–3] A number of studies have documented changes in the extent and duration of breastfeeding in developing countries (Table 1). Results are mixed, with some countries showing definite declines in either the extent or duration of breastfeeding, while others show increases in the same measures.

Breastfeeding declines do not necessarily follow rapid economic growth, as the examples of Taiwan,[4] Malaysia,[5,6] or South Korea[7] seem to suggest. In studies of Thailand,[8] Mexico,[9] and Panama,[10] economic growth during the time periods of analysis was considerable, yet the declines in breastfeeding were not as extensive as those observed in Taiwan or Malaysia (during the 1950–70 period).[4]

Patterns of decline in breastfeeding across population subgroups will not necessarily be uniform. In Taiwan (between 1967–68 and 1979–80), the extent and duration of breastfeeding decreased considerably for most socioeconomic groups,[3] but in the Dominican Republic the proportion of women who ever breast-fed actually increased in the urban areas while declining in the rural areas (from 1976 to 1983).[3]

For the Philippines, which has experienced only moderate socioeconomic change in the last decade, available studies present somewhat conflicting evidence about changes in breastfeeding behavior. While the country has been cited as one of those with large declines in breastfeeding,[11,12] the evidence on which the claim is based is derived from two noncomparable samples of women.[13,14] A recent review of a large number of small unrelated Filipino surveys concludes that there is little evidence either of a strong downward trend (since 1965) in the proportion of Filipino infants ever breast-fed, or of a significant decline (from 1960 to 1977) in the average duration of breastfeeding.* Zablan[15] also found only a minor decline in the extent of breastfeeding, and no change in breastfeeding duration.

TABLE 1—Summary of Selected Studies on Breastfeeding Trends in Developing Countries*

Country	Period of Analysis	Breastfeeding: Extent	Breastfeeding: Duration
Countries Which Experienced Rapid Economic Development			
Taiwan[1,3,4]	1967–68 to 1985	decrease	decrease
Malaysia[5,6]	1950 to 1970	decrease	decrease
	1970–74 to 1975–77	increase	constant
Korea[6]	1950 to 1970	constant	decrease
Singapore[3]	1951 to 1971	decrease	N/A
	1971 to 1980	increase	N/A
Countries with Lower Levels of Socioeconomic Development			
Thailand[9,10]	1969 to 1979	N/A**	decrease
	1975 to 1979	constant	N/A
	1979 to 1984	increase	increase
Mexico[11]	1976 to 1979	decrease	increase
Panama[12]	1976 to 1979	decrease	decrease
Jamaica[3]	1976 to 1983	increase	increase
Sri Lanka[3]	1975 to 1982	increase	constant
Guatemala	1978 to 1983	N/A	constant
Jordan[3]	1976 to 1983	N/A	constant, rural increase, urban
Dominican Republic[3]	1975 to 1983	increase, urban decrease, rural	increase

*Numbers in parentheses refer to sources.
**N/A = not available.

Methods

We analyzed data from the 1973 and 1983 National Demographic Surveys (NDS) of the Philippines, both of which had nationally representative samples of ever-married women, drawn by multistage stratified cluster sampling. Subsamples of 3,573 and 5,396 ever-married women and children from the 1973 and 1983 original samples, respectively, were used in the analysis. The results are based on Filipinos living in private households containing at least one ever-married woman at least 15 years old who had borne at least one child in the three years preceding each survey.

The original sampling weights assigned to the members of the subsamples are adjusted to preserve nationally representative patterns. Results on population composition, breastfeeding extent and duration all are based on the weighted subsamples.

The multiple-decrement life table procedure is used to compute the probability that a child is breast-fed at each age in months and thus the median duration of breastfeeding. The selected ages for which subgroup analyses are carried out are four, seven, and 13 months.

For both time periods there is strong heaping in the reports of cessation of breastfeeding at 12 months of infant age. This may reflect an actual tendency to stop at that age or an exaggeration because of digit preference in reporting. A small longitudinal study in the Philippines showed that the

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*Simpson-Hebert M: Infant feeding in the Philippines 1955–1983 and its effect on infant health: A literature review (unpublished manuscript). Quezon City, Philippines: Institute of Philippine culture, 1983.

proportion of infants[16] peaked at age 10–12 months. The same phenomenon was observed in a large scale longitudinal infant-feeding study of 3,080 infants conducted in the Cebu region from 1983 to 1985.**

Some of those who actually stopped breastfeeding at 10 to 11 months report a longer duration of one year, while some of those who actually stopped at 13 or 14 months report a shorter duration of one year. If the errors are spread symmetrically around the 12-month point, no bias in the estimated mean should result.

We chose to compare the median instead of mean duration, because the former is not likely to be as badly biased when there is digit heaping in reported duration.

Results

Breastfeeding, Socioeconomic and Demographic Trends

Changes in Population Composition—A number of selected demographic and socioeconomic indicators hypothesized to be associated with breastfeeding initiation,[17,18] suggest that from 1973 to 1983 the Philippine population not only became more urbanized, but also continued to concentrate in disproportionate numbers in the greater Manila area. Educational levels among ever-married females of ages 15 to 49 and their husbands rose substantially. There was a rise in the median level of real household income (at 1983 levels), and a doubling of the proportion of households with electricity. Occupations shifted from traditional agricultural work toward the unskilled sector, rather than toward the modern skilled sector of the work force.

Changes in household composition were the opposite of what one might expect in a modernizing society: the proportion of extended households increased by 4 per cent from 1973 to 1983, probably the result of rural-urban migration combined with the tendency of the immigrants to board with relatives or acquaintances; average household size declined with a slowly falling birth rate in the country (mean parity declined by about half a child, and contraceptive practice became more common).

Trends in the Extent of Breastfeeding

Figure 1 shows the likelihood of a child being breast-fed according to age by urban residency and by the two time periods of interest. The percentage shown as breast-fed at age 0 is the proportion "ever breast-fed." Table 2 presents information on the proportion of mothers who ever breast-fed classified by selected socioeconomic and demographic groups.

The data indicate a 5.4 per cent overall decline, over the 10-year period, in the percentage of infants ever breast-fed. This decline is consistent with that of Zablan[15] who reported about four percentage points decline from 1973 to 1978 and no decline from 1978 to 1983.

The downward trend is more pronounced in urban than rural areas. A larger decline in breastfeeding occurred among persons residing in Manila and Mindanao than among residents of other areas. Declines also occurred among women with 4–9 years of formal education and among women whose husbands had "mixed" or "traditional" occupations. Women in "modern" occupation or with at least 10 years of education increased breastfeeding initiation.

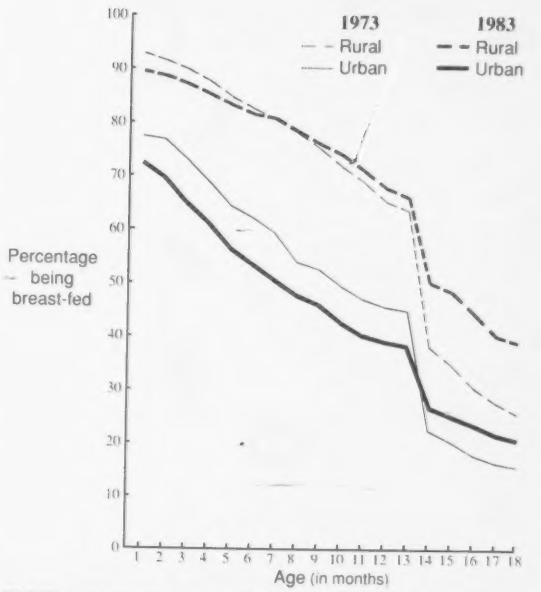

FIGURE 1—Percentage of Urban and Rural Children Breast-fed According to Age in Months, Philippines, 1973 and 1983.

Trends in the Median Duration of Breastfeeding

The median age of the infant at which breastfeeding ceased did not change from 1973 to 1983.*** Decrease in duration occurred among urban and Manila residents, women whose husbands had "modern" occupations, and women who were engaged in market activities. The decrease in these groups was due mainly to breastfeeding cessation in the first year of life. However, relatively large increases in the duration of lactation also occurred between 1973 and 1983 among rural residents, Visayan residents, women with 10 or more years of education, and women with traditional occupations. The increase for these subgroups occurred because in 1983 they were more likely to breastfeed their infants past the first year of life than were similar groups in 1973.

Discussion

We found that from 1973 to 1983 there was a decrease of 5.4 per cent in the extent of breastfeeding and essentially no change in median duration (12.4 in 1973 and 12.5 in 1983). Among women with "modern" occupations, there was an increase in the extent of breastfeeding but no change in median duration. This is not surprising since such women are better educated and better understand the value of breastfeeding, hence more likely to initiate nursing, yet also more likely to provide supplementary food at an early age and reduce duration of breastfeeding because of their occupation.

The decrease in breastfeeding initiation and decreases in probabilities of breastfeeding early in the first year of life are of particular concern because breastfeeding inhibits menstruation

**Cebu Study Team: Unpublished life table results of breast-feeding among Cebu infants, 1983–85. Chapel Hill: Carolina Population Center, University of North Carolina, 1987.

***Zablan (1983) reports an increase of one month of breast-feeding duration from 1978 to 1983; these differences in results may be attributed to the use of different subgroups of children as well as the difference in the technique of estimating duration.

TABLE 2—Incidence of Breastfeeding among Last Live Births Ages 0 to 36 Months at Time of Survey by Selected Characteristics, Philippines, 1973 and 1983 (percentage ever breast-fed, weighted life table results)[a]

Characteristics	Average Sample Size	Incidence		Unadjusted Difference[a]	95% Confidence Interval of Difference	
		1973	1983		Lower Limit	Upper Limit
Total Sample	4484	88.9	83.5	−5.4	−6.8	−4.0
Area of Residence						
Urban	1402	77.4	72.6	−4.8	−8.2	−1.4
Rural	3082	93.0	89.3	−3.7	−5.1	−2.3
Migration Status						
Urban Migrant	511	79.1	75.2	−3.9	−9.0	1.2
Urban Resident	678	75.4	72.3	−3.1	−8.1	2.0
Rural Resident	3082	93.0	89.3	−3.7	−5.1	−2.3
Region of Residence						
Manila	451	75.1	65.7	−9.4	−15.9	−2.9
Luzon (excluding Manila)	1936	88.2	85.0	−3.2	−5.3	−1.1
Visayas	1020	91.6	88.8	−2.8	−5.3	−0.3
Mindanao	1076	91.0	85.6	−5.4	−8.2	−2.6
Mother's Education (years)						
None	174	94.9	93.9	−1.0	−6.1	4.1
1–3	464	92.4	91.0	−1.4	−5.0	2.2
4–6	2142	92.6	89.1	−3.5	−5.2	−1.8
7–9	638	86.7	82.2	−4.5	−8.4	−0.6
10 ≅	1058	66.2	73.6	7.7	2.0	12.6
Father's Education (years)						
None	170	93.6	93.9	0.3	−5.2	5.8
1–3	538	94.4	91.6	−2.8	−5.8	0.2
4–6	1828	92.5	88.9	−3.6	−5.5	−1.7
7–9	452	89.5	84.9	−4.6	−9.0	0.2
10 ≅	1236	74.7	75.4	0.7	−3.0	4.4
Mother's Work Status						
Not Working	3463	90.9	85.4	−5.5	−7.0	−4.0
Presently Working	1021	82.6	76.6	−6.0	−9.4	−2.5
Works away from Home[b]	605	89.3	N/A	N/A	N/A	N/A
Works at Home	277	79.4	N/A	N/A	N/A	N/A
Mother's Occupation						
Modern	219	53.0	63.0	10.0	0.1	19.9
Transitional	76	84.6	73.4	−11.2	−24.6	2.2
Mixed	454	82.0	78.8	−3.2	−8.5	2.1
Traditional	269	95.9	92.3	−3.6	−7.7	0.5
Father's Occupation						
Modern	266	69.4	67.4	−2.0	−10.0	6.0
Transitional	268	78.1	76.4	−1.7	−8.9	5.5
Mixed	1280	84.2	78.0	−6.2	−9.4	−3.0
Traditional	2247	94.6	92.2	−2.4	−3.8	−1.0
Income Level[c]						
Low	1959	93.9	91.3	−2.6	−4.3	−0.9
Middle	1479	89.4	82.4	−7.0	−1.5	−4.5
High	892	75.1	69.7	−5.4	−9.9	−0.9

[a]The differences for the total sample between 1973 and 1983 adjusted to keep sample characteristic as equal for the two periods, (i.e., for example to adjust the 1983 urban-rural mix to be the same as that) 1973 are: −4.0, −4.1, −2.1, −2.6, −5.7, −4.6, −3.0 and −4.6 when adjustment is made for area of residence, region, mother's education, father's education, mother's working status, mother's occupation and father's occupation. The 1973 sample composition is used as reference population for the 1983 adjustment.

[b]Information on the location of woman's work is unavailable in 1983.

[c]In 1973, the income levels were based on the father's income while in 1983 the grouping is based on household income.

to some extent, and the protective effects of breastfeeding against morbidity are most pronounced during infancy.[19,20] In a country like the Philippines where health care and contraceptive utilization are limited, decreased breastfeeding initiation may have a deleterious public health impact.

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Johnson Foundation Funds Model Program for Elderly in Housing Projects

The Robert Wood Johnson Foundation has awarded grants to 10 state housing finance agencies (HFAs) to develop new ways to finance and deliver support services for frail older people living in federally subsidized private housing for the elderly. Under its Supportive Services Program in Senior Housing, the Princeton-based Foundation has committed up to $4 million over a three-year period to support the efforts of the 10 HFAs. The foundation will provide up to $400,000 to each HFA over the program's duration.

The grants will enable HFAs to work with local housing projects to provide the elderly with needed services such as: arranging emergency medical alert systems, supervising medication schedules, providing transportation, and assisting residents with difficult household chores. The goal of the program is to establish models of service delivery that will enable elderly tenants in subsidized housing to remain independent.

The grant awards will be distributed in two successive phases. An initial one-year grant will fund a planning effort to be reviewed before the two-year implementation grant is approved. During the planning phase, the HFAs will:

- identify housing projects with a significant proportion of people in need of support services;
- assess the training needs of the housing projects' management and staff; and
- develop the financing mechanisms necessary to implement services.

In the implementation phase, HFAs will use their own reserve funds as well as those of the housing projects to establish relationships with service providers. It is expected that a combination of client fees and housing-based reserve subsidies will pay for services requested by the residents.

Planning grants under the first phase of the program were awarded to the following HFAs:

- $144,807 to the Colorado Housing and Financing Authority, Denver;
- $49,598 to the Illinois Housing Development Authority, Chicago;
- $88,517 to the Maine State Housing Authority, Augusta;
- $204,000 to the Massachusetts Housing Finance Agency, Boston;
- $62,697 to the New Hampshire Housing Finance Authority, Manchester;
- $182,935 to the New Jersey Housing and Mortgage Finance Agency, Trenton;
- $179,874 to the Pennsylvania Housing Finance Agency, Harrisburg;
- $107,669 to the Rhode Island Housing and Mortgage Finance Corp., Providence;
- $64,786 to the Vermont Housing Finance Agency, Burlington; and
- $71,620 to Virginia Housing Development Authority, Richmond.

The program is directed by James J. Callahan, Jr., PhD, who also directs the Policy Center on Aging at Brandeis University in Waltham, Mass. The Robert Wood Johnson Foundation was established as a national philanthropy in 1972. Since then, it has awarded more than $906 million in grants to improve health care in the United States. For further information, contact the RWJ Foundation, P.O. Box 2316, Princeton, NJ 08543-2316. Tel: (609) 452-8701.

A Recent Increase of Breastfeeding Duration in Jakarta, Indonesia

M. Riduan Joesoef, MD, PhD, Joseph L. Annest, PhD, and B. Utomo, MD, MPH

Abstract: Comparison of breastfeeding practices from two similar surveys conducted in Jakarta, Indonesia shows an increase in breastfeeding duration from 14.4 months in 1976 to 19.8 months in 1983. This increase was predominantly among noneducated women (20.3 months among noneducated women versus 0.9 months among women with greater than elementary education). Among working women breastfeeding duration declined slightly (1.1 months). Unlike industrialized countries where increases in breastfeeding have occurred initially among the higher socioeconomic groups, recent increases in Jakarta have occurred initially among the lower socioeconomic groups. If breastfeeding practices among working or better educated women who delivered mostly in hospitals or clinics are to be improved, the breastfeeding promotion programs should be targeted to these groups. (*Am J Public Health* 1989; 79:36–38.)

Introduction

Until recently, breastfeeding has been declining in most developing countries, especially in cities and peripheral urban areas.[1–4] The decline in breastfeeding practices has had a detrimental impact on infant health and survival.[5] As a result, many developing countries where infant health and survival are a major health problem have launched national programs to promote breastfeeding.[6]

Breastfeeding practices are usually examined by a single cross-sectional survey at one point in time. Obtaining equivalent surveys on breastfeeding practices at two points in time is difficult, particularly in cities where breastfeeding patterns are changing most rapidly. Recently, however, two studies—one in Malaysia,[7] the other in the City of Sao Paulo, Brazil[8]—evaluated breastfeeding trends and found a revival of breastfeeding practices.

In Indonesia the National Commission on Breastfeeding, a nongovernmental body, was founded in 1977 to coordinate a national program to promote breastfeeding. Two similar household surveys conducted in 1976 and 1983 provided the opportunity to examine changes in breastfeeding practices. In this report, we examined changes in breastfeeding practices in Jakarta, a city with a population of more than 6.5 million in 1980.

Methods

We used data from the two large household surveys conducted in 1976 (Indonesia Fertility Survey) and in 1983 (Indonesia Contraceptive Prevalence Survey). Because only Jakarta was covered by both surveys, we restricted our comparison of breastfeeding practices to Jakarta. Both surveys were based on a probability sample designed to be approximately self-weighting; clusters of households were randomly selected, and in each household all ever-married women aged less than 50 years were enumerated. The principal reports of the two surveys contain a detailed methodological description.[9,10]

In both surveys, a similar questionnaire design was used in which interviewers asked questions about the initiation and duration of breastfeeding and whether the child was currently breastfeeding. In the 1983 survey, breastfeeding questions were asked only to woman whose last living child was 3 years of age or younger. These breastfeeding questions referred only to the woman's last (most recent) live birth, not to her "next-to-last" live birth or other births, i.e., breastfeeding in her current birth interval.[11] In the 1976 survey, breastfeeding referred to the woman's last (most recent) live birth, regardless of the child age. To make the two surveys comparable in terms of breastfeeding definition, for the 1976 survey we restricted our analysis to the women whose last living child aged 3 years or younger at the time of the survey.

In this study, breastfeeding refers to the period from birth to complete cessation of breastfeeding, regardless of supplemental foods. Life-table techniques were used to calculate the median duration of breastfeeding and the proportion of women who were still breastfeeding their last living child at various durations of postpartum.[12] We used SPSS[x] (Survival Procedure) software package to construct the life-table.[13] The following variables and categories were examined: mother's work status at the time of the survey (yes, no); mother's and father's education (no school, ≤ elementary, > elementary); number of living children (< 3, ≥ 3); mother's age (< 30 years, ≥ 30 years); and desire for more children (yes, no).

Results

The total percentage of women who initiated breastfeeding was high and stable for both surveys—98 per cent in 1983 and 93 per cent in 1976 (Table 1). The rates of initiation were high and slightly higher in the 1983 survey across all factors examined (Table 1).

In both surveys the proportion of women still breastfeeding during the first month of postpartum was also high (0.96 in 1983 and 0.98 in 1976). After the first month, the proportion of women breastfeeding declined in a similar pattern up to six months for both surveys (Figure 1). After six months, the proportion of women breastfeeding declined less rapidly in 1983 than in 1976.

Overall, the median duration of breastfeeding increased substantially from 14.4 months in 1976 to 19.8 months in 1983 (Table 2). The change in breastfeeding duration varied across sociodemographic groups with the lowest educated group having the largest increase (Table 2). For mothers with no school education the median duration increased substantially from 15.7 months in 1976 to 36.0 months in 1983, but for mothers with education greater than elementary the median duration increased only from 13.1 months to 14.0 months. A similar pattern was seen for father's education. For mothers who did not work at the time of the survey, the median duration increased from 13.8 months to 20.3 months, and for

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TABLE 1—Percentage of Women Initiated Breastfeeding, by Survey for Various Characteristics, Jakarta, 1976, 1983

	Surveys			
	1983		1976	
	% (N)	SE	% (N)	SE
Mother's education				
No School	98 (91)	1.5	95 (242)	1.4
≤ elementary	98 (532)	0.6	92 (208)	1.9
> elementary	98 (364)	0.7	93 (158)	2.0
Father's education				
No School	100 (29)	0	89 (50)	4.4
≤ elementary	98 (352)	0.7	94 (266)	1.5
> elementary	97 (606)	0.7	94 (292)	1.4
Mother's current work status				
Working	97 (135)	1.5	92 (161)	2.1
Not working	98 (852)	0.5	94 (447)	1.1
Number of living children				
0–2	98 (486)	0.6	93 (264)	1.6
3+	98 (501)	0.6	93 (344)	1.4
Mother's age				
< 30 years	99 (675)	0.4	93 (396)	1.3
≥ 30 years	96 (312)	1.1	94 (212)	1.6
Desire for more children				
Yes	98 (431)	0.7	92 (280)	1.6
No	98 (490)	0.6	95 (211)	1.5
Total	98 (1009)	0.4	93 (608)	1.0

TABLE 2—Median Duration of Breastfeeding, by Survey for Various Characteristics, Jakarta, 1976, 1983

	Surveys		Difference (in months)
	1983	1976	
Mother's education			
No School	36.0	15.7	20.3
≤ elementary	20.9	13.8	7.1
> elementary	14.0	13.1	0.9
Father's education			
No School	36.0	19.5	16.5
≤ elementary	24.3	14.6	9.7
> elementary	16.5	12.7	3.8
Mother's current work status			
Working	15.3	16.4	−1.1
Not working	20.3	13.8	6.5
Number of living children			
0–2	18.1	13.9	4.2
3+	25.0	14.8	10.2
Mother's age			
< 30 years	19.2	12.8	6.4
≥ 30 years	24.5	17.4	7.1
Desire for more children			
Yes	18.8	12.0	6.8
No	20.9	16.2	4.7
Total	19.8	14.4	5.4

mothers who worked, the duration decreased slightly from 16.4 months to 15.3 months (Table 2).

Discussion

In this investigation, we found an increase in breastfeeding duration in Jakarta during the study period from 1976 to 1983. The increase was the highest among noneducated women. Essentially no changes were seen for women with greater than elementary education and for working women. The rates of breastfeeding initiation were practically unchanged across various sociodemographic groups in both surveys. Consistent with previous studies, the rates of breastfeeding initiation and duration of breastfeeding were high.[14–16] Bracher and Santon[14] who studied the breastfeeding practices in Central Java between 1976 to 1978 found that "fully two-fifths of children can be expected to be unweaned by the longest duration recorded in this study" (29 months).

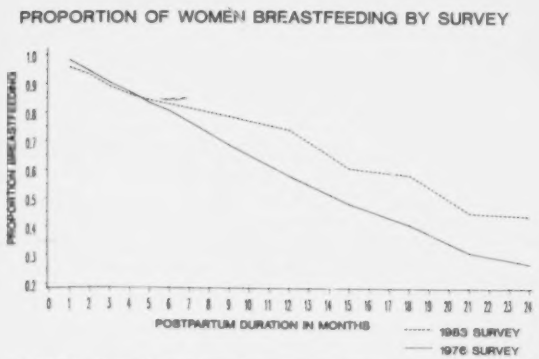

FIGURE 1—Proportion of Women Breastfeeding by Survey, Jakarta, Indonesia, 1976 and 1983

In another study in 1973, Djumadias and Mastaitella found that in South Sulawesi (one of the provinces in Indonesia) 38 per cent of the women breastfed their children between two and three years.[16] These total breastfeeding durations are in agreement with the total breastfeeding duration of the present study.

The increase in breastfeeding duration is not explained by differences in the study population, sampling, or questionnaire design. The sociodemographic characteristics of the two study populations, sampling, and questionnaire design for the two surveys were similar. Some effect of recall bias, however, or some other form of nonsampling error cannot be ruled out. Because the degree of the recall bias was proportional to the length of time the women had to recall the information, limiting our analysis to breastfeeding information from the last live child who was born three years before interview would tend to minimize the recall bias. As a secondary analysis, we evaluated breastfeeding duration only for mothers who were still breastfeeding at the time of interview (those least affected by recall bias); findings were similar. In both surveys we found insignificant, similar heaping of breastfeeding duration at 12, 18, and 24 months postpartum. Thus, we believe that the increase in breastfeeding duration is not an artifact of reporting bias, but a true reflection of a change in breastfeeding duration. These changes may apply to other major Indonesia cities (cities that were surveyed in 1983) because the breastfeeding patterns for these cities and Jakarta were similar.[17]

The vigorous efforts by the National Family Planning Coordinating Board to reduce fertility also might have had an impact on differences in breastfeeding duration between the two surveys. McNicoll and Singarimbun stated that modern contraception was responsible for fertility reduction in Indonesia.[18] Bracher and Santon who studied the breastfeeding practices in Central Java between 1976 to 1978 found that two-fifths of the mothers who weaned their children were known to be pregnant.[14] Fertility reduction might prolong breastfeeding duration because mothers who were breastfeeding might have weaned their children when they became

pregnant. Comparison of the rates of current pregnancy between the two surveys supports this hypothesis. The rate was 15.7 per cent in 1976 and 9.3 per cent in 1983.

The national breastfeeding promotion programs, which were started in 1977, might have had an impact on breastfeeding duration. However, we believe that these programs have failed to convey the correct messages or have inadequately reached the working or better educated women who delivered mostly in hospitals or clinics.[17] A recent survey of rooming-in and breastfeeding management in teaching hospitals of Jakarta and other major Indonesian cities supports our belief.[19] Only 50 per cent of mothers who delivered their babies in these hospitals chose "rooming-in" with the babies. No mothers and very few health care providers mentioned frequent suckling, relaxation, and emotional support as important determinants for prolonged breastfeeding. In addition, this hospital survey found three of 15 hospitals with posters promoting infant formula, and at least one hospital administrator in each of the cities in the hospital survey admitted that mothers in their hospital did receive bottle-milk samples.

According to the World Health Organization proposed model of breastfeeding practices, the elite, better educated people will lead the breastfeeding revival, followed by the poor urban people as has been observed in the United States and Sao Paulo, Brazil.[5,8] Contrary to this model, in Jakarta, the less educated women led the breastfeeding revival. These findings are similar to those found in the study conducted in the neighboring country of Malaysia, where cultural and religious values are similar to those in Indonesia.[7] The increase in Malaysia, however, was not in the duration but in the initiation of breastfeeding.

ACKNOWLEDGMENTS

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Patient Education Conference on Arthritis Scheduled for March 1989

The National Arthritis Advisory Board and the National Institute of Arthritis and Musculoskeletal and Skin Diseases will sponsor a conference, "Financing Patient Education in Arthritis Musculoskeletal and Skin Diseases," March 23–24, 1989 in Atlanta, Georgia. The conference will address such topics as the cost-effectiveness of patient education in chronic disease, assuring the quality of patient education, and various perspectives on financing and third-party reimbursement of patient education. In addition, ongoing programs that might serve as models will be presented, and participants will meet in small groups to discuss future activities and plans that might be undertaken on a regional basis.

The goals of the conference are:

- To present what is known about the clinical and fiscal efficacy of patient education in arthritis and musculoskeletal and skin diseases,
- To document what is known and identify areas requiring further investigation, and
- To stimulate interest in obtaining third-party reimbursement for patient education in arthritis and musculoskeletal and skin diseases.

For more information, contact:

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Cryptosporidiosis and Surface Water

Margaret M. Gallaher, MD, Joy L. Herndon, MS, Linda Jean Nims, MS, Charles R. Sterling, PhD, Debra J. Grabowski, MS, RS, and Harry F. Hull, MD

Abstract: In the period July through October, 1986, 78 laboratory-confirmed cases of cryptosporidiosis were identified in New Mexico. To determine possible risk factors for development of this disease, we conducted a case-control study: 24 case-patients and 46 neighborhood controls were interviewed. Seventeen (71 per cent) of the 24 case-patients were females, seven (29%) were males; their ages ranged from 4 months to 44 years, median 3 years. There was a strong association between drinking surface water and illness: five of the 24 case-patients, but none of the 46 controls drank untreated surface water. Among children, illness was also associated with attending a day care center where other children were ill (odds ratio = 13.1). (*Am J Public Health* 1989; 79:39–42.)

Introduction

Cryptosporidium, first recognized as a cause of diarrhea in humans in 1976,[1] has become known as a common cause of diarrhea worldwide.[2,3] No effective therapy is known; the diarrhea is often prolonged and accompanied by weight loss. In immunocompromised persons, cryptosporidiosis can cause a life-threatening dehydration.

Cryptosporidium can be transmitted through contact with animals, especially cattle and sheep,[4] through person-to-person contact, especially in day care centers,[5] and through a contaminated community water source.[6] In several reported outbreaks, the source of the infection was unknown.[7,8]

In 1984, 24 cases of cryptosporidiosis were reported to the New Mexico Office of Epidemiology, 13 of them were associated with two day care centers. In 1985, 24 cases were reported. In July 1986, a cluster of 20 cases of cryptosporidiosis was recognized in Bernalillo County (the city of Albuquerque). Because these cases did not appear to be related and the source of infection was unclear, a case-control study to evaluate potential risk factors for illness was conducted.

Methods

Since February 1984, all stool samples submitted to the New Mexico Health and Environment Department's Scientific Laboratory Division for ova and parasite testing have been examined for *Cryptosporidium*. Stool samples are concentrated using a formalin-ethyl acetate technique. A heat-fixed fecal smear is stained with Kinyoun modified acid-fast stain[9] and counterstained with Loeffler methylene blue. The specimen is then examined under a light microscope for the presence of *Cryptosporidium* oocysts.

In August 1986, the Office of Epidemiology sent a letter to primary care physicians in Bernalillo County, alerting them to the cluster of cases, reviewing the symptoms and diagnosis of cryptosporidiosis, and requesting that stool samples be sent if cryptosporidiosis was suspected.

A case-patient was defined as a Bernalillo County resident with laboratory-confirmed cryptosporidiosis reported to the Office of Epidemiology between July 1 and October 1, 1986. If more than one person with laboratory-confirmed cryptosporidiosis was identified within a household or day care group, only the one with the earliest onset of symptoms was included in the study.

Controls were matched for age, sex, and neighborhood using a reverse telephone directory. The home of the case-patient was identified in the directory. The homes listed on either side of the case-patient were contacted by telephone alternately until suitable controls were identified. Case-patients 10 years old or younger were matched within two years of age, and those over 10 years were matched within five years of age. A control could not have had diarrhea in the four weeks before the onset of disease in the case-patient to which he/she was matched. Twenty-two of the 24 index cases were matched with two controls each. Two index cases were matched with only one control.

A standard questionnaire, administered via telephone to both case-patients and controls, requested information on size of household, day care attendance or contact, travel, surface water exposure, pets and domestic animal exposure during the four weeks before onset of illness in the case-patient.

Two surface water samples were collected on October 27, 1986 from a lake in southern New Mexico and a lake in Arizona where three of the case-patients had been exposed to surface water. These samples were tested for *Cryptosporidium* and *Giardia* using a fluorescent labeled monoclonal antibody technique.[10,11]

Using the matched triplets or pairs, we calculated Mantel-Haentzel chi-square, odds ratios (OR), and 95% confidence intervals (CI) and fitted a conditional logistic regression to determine which variables were independently associated with infection with *Cryptosporidium*.

Results

Of the 56 laboratory-confirmed cases of cryptosporidiosis in Bernalillo County reported between July 1 and October 1, 26 met the definition of a case-patient. Two (8 percent) of the cases were lost to follow up (a 1-year-old female who attended a day care center and a 43-year-old female). Of the 30 cases excluded from the study, two were patients with acquired immune deficiency syndrome who died during the course of the study, 10 were household

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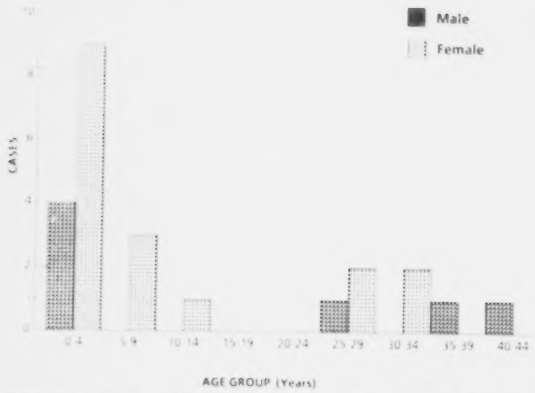

FIGURE 1—Cryptosporidiosis, by Age Group and Sex, New Mexico, 1986

contacts of case-patients, and 18 were day care contacts of case-patients.

The 24 remaining case-patients ranged in age from 4 months to 44 years with a median age of 3 years. Figure 1 shows the age and sex distribution of the index cases. Seventeen of the cases were 10 years of age or younger; 13 of whom were females; 17 of all 24 cases were female. Thirteen (54 per cent) of the case-patients were White and 11 (46 per cent) were Hispanic. Controls showed similar demographic characteristics.

Case-patients had onset of symptoms from May 28 to September 2, 1986 (Figure 2). The duration of symptoms ranged from five to 60 days with a median of 21 days. The most frequent symptom (Table 1) was diarrhea followed by flatulence and abdominal pain.

Matched set analysis identified two possible sources of infection: surface water (rivers, lakes and streams) and day care centers (Table 2). Five case-patients drank untreated surface water, two in Mexico, and three in southern Colorado and New Mexico; none of the controls drank surface water. Case-patients reported more swimming in surface water in

TABLE 1—Frequency of Reported Symptoms of Cryptosporidiosis, New Mexico, 1986 (N = 24)

Symptom	No.	%
Diarrhea	23	96
Flatulence	20	79
Abdominal Pain	17	67
Nausea	15	58
Fever	13	54
Vomiting	12	46
Abdominal distention	12	46

the four weeks prior to illness. The time between exposure to surface water (swimming or drinking) and onset of illness ranged from four to 21 days with a median of seven days. Fourteen (82 per cent) of 17 household members with exposures to surface water similar to the case-patients also became ill with diarrhea. These household members became ill within two to seven days of the onset of illness in the case-patient.

Case-patients were substantially more likely to attend a day care center where other children were ill with diarrhea than were their matched controls (Table 2). Attendance at day care centers where no other children were ill was not related to risk.

The logistic regression model for multivariate analysis included all variables except drinking surface water, which did not vary. Infection was associated with swimming in surface water, attending a day care center, camping, and having a pet which was less than six months of age or was ill, with odds ratios all greater than 3.0. Having a pet over six months of age was associated with a decreased risk of disease. All of these estimates were unstable, with 95% confidence limits including 1.0.

Among the 24 case-patients, eight were exposed to surface water, eight were day care center attendees or employees, six were household contacts of a day care center attendee, and eight had young and/or ill pets. Four had none of these exposures.

The surface water samples from New Mexico and Arizona where cases had been exposed did not yield *Cryptosporidium* or *Giardia* on testing.

Two of the 24 index cases were coinfected with *Giardia*.

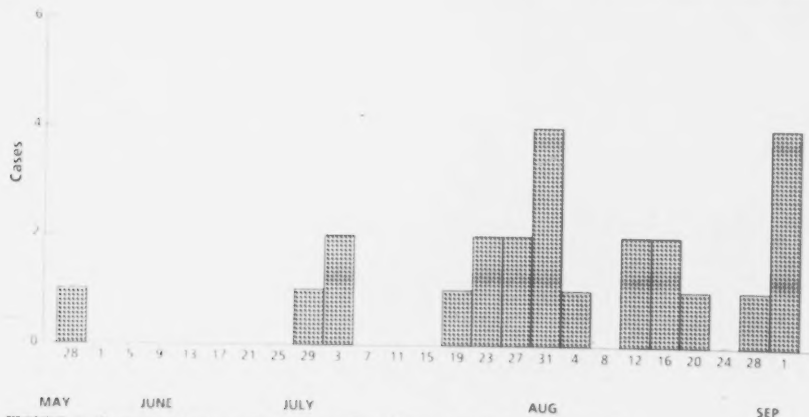

FIGURE 2—Cryptosporidiosis, by Date of Onset, New Mexico, May 28–September 4, 1986

TABLE 2—Exposures Associated with Cryptosporidiosis

Exposure	No. of Cases Exposed	No. of controls exposed			Odds Ratio	95% Confidence Interval
		0	1	2		
Drink surface water	0	17 (2)	0	0	00	2.44, 00
	1	5	0	0		
Swim in surface water	0	14	3	0	3.7	1.02, 13.5
	1	4 (2)	1	0		
Attend a day care center where other children were ill	0	15 (1)	1	0	13.1	2.58, 66.4
	1	6 (1)	0	0		
Household contact of day care center attendee	0	12 (1)	4	0	3.0	0.91, 9.85
	1	5 (1)	1	0		

Numbers in parentheses represent cases with only one control.

One of them attended a day care center. The other had no apparent risk factors for cryptosporidiosis or giardiasis.

Discussion

Person-to-person transmission of cryptosporidiosis is well recognized, especially in day care centers.[5,12] Lack of hygiene and frequent mouthing of objects among diapered babies and toddlers can lead to the rapid spread of this and other enteric infections in day care centers and homes.[13] This study confirmed the association between attending a day care center, especially where other children are ill, and cryptosporidiosis.

Cryptosporidiosis has also been found to be a cause of traveler's diarrhea, particularly in people traveling to less developed countries.[2,14] In this study, two (8 per cent) of the index cases had recently traveled to Mexico; both of these case-patients also drank untreated surface water in Mexico.

Although puppies and kittens have been shown to excrete *Cryptosporidium*,[4] these pets have not previously been demonstrated to transmit cryptosporidiosis to humans. Contact with a young and/or ill pet was associated with an increased risk of disease in this study, but the numbers are too small to be convincing. Nevertheless, this observation is in keeping with what is known of cryptosporidiosis in domestic animals: young animals contract cryptosporidiosis more often than older animals, and the severity and duration of illness is inversely related to the age of the animal.[1,15]

This study suggests for the first time that drinking surface water may be a mode of transmission of cryptosporidiosis in the United States. This is feasible since waterborne cryptosporidiosis has been previously demonstrated in a community in Texas due to a contaminated artesian well.[6] The plausibility of waterborne transmission is increased by the time relationship between exposure to surface water and onset of illness. The median incubation period of seven days in those case-patients exposed to surface water is consistent with what is known of the incubation period for cryptosporidiosis.[16]

Cryptosporidium sp. has been found in a broad variety of both domestic and wild animals. Cattle and sheep appear to be important hosts, but mice, snakes, rabbits, and deer have also been found to excrete the organism.[4,15] Much of the recreational water in New Mexico and other western states drains land which has multiple uses including the grazing of livestock. Livestock could contaminate surface water with *Cryptosporidium* by direct deposit of feces in the water or via surface runoff washing animal feces into the water. Similarly, contamination of the water by wild animal feces might also play an important role. New Mexico received heavier than normal rainfall during the summer months of 1986, a factor which may have promoted contamination of surface water.

The finding of an elevated risk of cryptosporidiosis among persons who had recreational contact with, but did not drink surface water, supports waterborne transmission of the disease. Recreational surface water contact, specifically swimming, has been shown to be a risk factor for other enteric infections.[17,18] Swimmers become infected by swallowing small amounts of contaminated water, especially following submersion. The infective dose of *Cryptosporidium* is unknown, but is probably small, because of the organism's ability to complete its life cycle in a single host and to reinitiate the developmental cycle in that host.[4] A small infective dose for cryptosporidiosis would facilitate transmission by water contact without frank water consumption.

Our studies of surface water in New Mexico were limited in scope, but failed to demonstrate the presence of *Cryptosporidium*. These studies were done two months after the outbreak and probably represent a failure to detect intermittent, low-level, or seasonal contamination. Additionally, the samples were taken from a large lake and it was difficult to determine the exact site of exposure. Studies of surface water conducted in Arizona showed that there was widespread contamination of streams with *Cryptosporidium* during the summer of 1986,[11] but that most of the water sources did not contain *Cryptosporidium* during the fall months.*

Cryptosporidiosis occurs more commonly in summer and fall in these and other data.[3,4] Increased outdoor activity during the summer and early fall with drinking of or swimming in surface water may partially explain the seasonal variation.

Coinfection with *Giardia lamblia* has been found in 2 per cent to 23 per cent of cryptosporidiosis cases in previously described outbreaks[19–21] and 8 per cent in this study. The coincidental occurrence of two parasitic infections suggests a common mode of transmission. Person-to-person transmission of both giardiasis and cryptosporidiosis in day care centers has been well described.[12] Waterborne transmission of giardiasis from surface water is also well known.[22,23] Our

* Unpublished data, CR Sterling, University of Arizona, 1986.

study suggests that waterborne transmission of cryptosporidiosis may explain coinfection with these two parasites, even though neither of the coinfections in this study occurred in persons who had exposure to surface water.

The median duration of diarrhea in the study population was longer than the three to 14 days described by others,[1,17,25] perhaps because patients who had stool samples submitted for ova and parasite examination had diarrhea that lasted longer than three to five days and did not respond to usual symptomatic treatment.

The predominance of young children found here has been seen in other studies, and follows the same pattern seen in domestic animals. In addition, young children are more likely to come to medical attention earlier in the course of a diarrheal illness than adults and more likely to be tested for the cause of their diarrhea. The explanation for the predominance of females seen in this group of children is not clear and may reflect only incomplete case ascertainment.

Further study of the role of surface water in transmission of cryptosporidiosis is clearly needed. Meanwhile, boiling of surface water rather than chemical disinfection may be prudent, since *Cryptosporidium* is resistant to most commonly used chemical disinfectants including chlorine and iodine,[15,24] but oocysts are destroyed by heating above 65°C for 30 minutes.[24] Boiling also destroys other organisms, including *Giardia*, that may be in the water. We also recommend that physicians consider cryptosporidiosis in the differential diagnosis of persons with diarrhea and a history of surface water contact.

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New Report Cites Rise in AIDS Funding by Foundations

A new report by The Foundation Center cites a significant increase in private foundation support for AIDS/HIV programs in the past year, with grants awarded by an increasingly diverse group of funders supporting an ever-wider range of programs. *AIDS Funding: A Guide to Giving by Foundations and Charitable Organizations* reports that, of the $51.6 million awarded by more than 150 private foundations since 1983 (including 40 of the top 100 foundations), $31 million has been provided in the period between August 1987–August 1988.

Roughly half of the funds awarded to date have been made by the Robert Wood Johnson Foundation of New Jersey. Only 3 per cent of all larger foundations—those with $1 million or more in assets or $100,000 or more in total giving (roughly 5,000 of the 25,000 foundations in the United States—have as yet awarded AIDS grants.

While 43 private foundations have awarded cumulative total grants of at least $100,000, only six foundations have awarded, in numbers, 15 or more grants for AIDS/HIV. Five foundations now include AIDS-specific funding as a priority interest.

The Foundation Center is located at 79 Fifth Avenue/16th Street, New York, NY 10003. Tel: (212) 620-4230.

A Prospective Follow-Up Study of Cancer Mortality in Relation to Serum DDT

HARLAND AUSTIN, DSc, JULIAN E. KEIL, PhD, AND PHILIP COLE, MD, DrPH

Abstract: Serum DDT and DDE levels were measured in 919 subjects in 1974 and 1975. Two-hundred and nine of the subjects died, including 54 from cancer, during a 10-year prospective follow-up period. There was no relation between either overall mortality or cancer mortality and increasing serum DDT levels. There was weak evidence of a positive relation between respiratory cancer mortality and serum DDT. The literature on DDT and human cancer is reviewed, and it is concluded that the evidence does not support the opinion that DDT is a human carcinogen. (*Am J Public Health* 1989; 79:43–46.)

Introduction

The widespread agricultural use of dichloro-diphenyl-trichloroethane (DDT) in the United States began in about 1945 and increased steadily until about 1960. At that time its use declined slightly because of increased resistance of pests to it. In June 1972, the United States Environmental Protection Agency (EPA) banned all crop uses of DDT. Outside the United States DDT is still used in agriculture, mainly for cotton crops, and for the control of insect-borne diseases, particularly malaria, in tropical countries.[1]

A major reason for the EPA ban of DDT was that it caused cancer in animals. DDT has been reported to induce liver tumors, lymphomas, and lung tumors in mice and liver tumors in some experiments with rats. On the other hand, feeding studies in hamsters were judged as negative; while those with dogs and monkeys are considered inconclusive.[2–4]

Because of recent litigation regarding putative exposure to DDT resulting from environmental contamination in Triana and other parts of northern Alabama, there is considerable public concern regarding the potential carcinogenicity of DDT in humans. There have been few epidemiologic studies of DDT and cancer, but some have reported an excess of lung cancer mortality among men occupationally exposed to DDT.[5,6] This analysis was undertaken, in part, because of this public concern and, in part, because of the suspected association between lung cancer and DDT.

DDT is metabolized very slowly by human beings, so that DDT metabolites are detectable in the fat and sera of people long after exposures end. This paper describes the overall mortality and the cancer mortality experience of a cohort of people whose serum DDT levels were measured in the mid-1970s and who were then followed prospectively for 10 years.

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Abbreviations used:
DDT = dichloro-diphenyl-trichloroethane
EPA = Environmental Protection Agency
PH = proportional hazards
RR = relative rate
SMR = standardized mortality ratio(s)
CI = confidence intervals

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Methods

The study has been described in detail elsewhere.[7,8] In 1960, 2,283 adult residents of Charleston, South Carolina were enrolled in a prospective follow-up study. In 1974 and 1975 an attempt was made to obtain a venous blood sample from the 1,708 survivors of the original cohort (468 White men; 602 White women; 310 Black men, 328 Black women). Blood specimens were obtained for 304 White men, 327 White women, 204 Black men, and 84 Black women and analyzed for p,p′ DDT and p,p′ DDE by gas liquid chromatography using the Dale-Cueto modified method.[9,10] These 919 subjects comprise the cohort reported on here.

The subjects were recalled between 1984 and 1985 by which time 209 (23 per cent) had died. Seven-hundred of the subjects were determined to be alive, while the vital status of 10 was unknown (99 per cent follow-up rate). The cause of death among decedents was classified in two ways. First, the cause of death as listed on the death certificate was coded according to the 9th Revision of the International Classification of Diseases. Additionally, all deaths were reviewed by a medical committee and a cause of death was assigned by them on the basis of all available information. The determination of the medical committee is used to classify cause of death for internal comparisons comparing one group of study subjects with another. To compare study subjects with an external population, the encoded cause of death is used. Discrepancies between the two classifications account for slight differences in the number of cancer deaths in the internal and external analyses.

Statistical Procedures

Total serum DDT among study subjects was estimated by multiplying serum p,p′ DDE (in parts per billion) by 1.114 and adding to this serum, p,p′ DDT (ppb). This accommodated the slight reduction in the molecular weight of DDE after its metabolism from DDT. Since serum p,p′ DDT and p,p′ DDE comprise about 95 per cent of all serum DDT metabolites and isomers,[11] this combination of the two approximates total serum DDT. The distribution of serum DDT was examined among the 919 subjects and divided into terciles. The relative mortality rate (RR) was obtained for the second and third tercile using the lowest tercile as the referent category.

In the internal comparisons below, the relation between serum DDT and mortality is evaluated after adjustment for age, race, gender, years of schooling, and cigarette habit (non-smokers, ex-smokers, and current smokers) through the use of Cox's proportional hazards (PH) model.[12] Indicator variables were used to denote the terciles of DDT. The antilogarithm of the parameters from such a model is inter-

pretable as the RR of the cause of death of interest for persons in the second or third tercile compared with those in the first. Dose-response was evaluated by including the logarithm of serum DDT in a PH model or by assigning scores of 0, 1, and 2 for the first, second, and third terciles, respectively, and including this ordinal exposure variable in a PH model. The logarithm (base 10) of serum DDT was used as the continuous exposure since the distribution of serum DDT is asymmetric. The 10 subjects whose vital status is unknown were considered lost to follow-up at the midpoint of the interval between the date their blood was assayed and January 15, 1985.

As an additional test for association between serum DDT and mortality, the mean of the logarithm of serum DDT was compared for those who had died from all causes or from some specific cause with those who had not. These means were obtained from a multiple linear regression model which included terms for the potential confounders, an indicator variable denoting whether or not the individual had died of that cause, and with the logarithm of serum DDT as the outcome variable.

In addition to the internal analyses described above, the mortality experience of study subjects was compared to that of residents of the United States or of South Carolina. The age- (six 10-year groupings), gender-, and race-specific person-years of observation of subjects were obtained and the corresponding age-gender-race specific cancer mortality rates of the 1980 United States or South Carolina population were applied to those person-years to obtain an expected number of deaths.[13] The observed number of deaths divided by the expected number is the standardized mortality ratio (SMR). The Poisson distribution was used for the purposes of statistical testing and estimation.[14] All P-values in the results section are two-tailed.

Results

The median age of study subjects at the beginning of the follow-up period was 60 years (range: 39 to 89 years). The mean serum DDT level was 48 ppb with a standard deviation of 36 ppb (on the logarithm scale the mean is 1.59 and the standard deviation is 0.30). The mean level of serum DDT was about 6 per cent higher among men than women, and about 14 per cent higher among Blacks than Whites. Serum DDT levels increased with age and were lower among the more educated. The mean DDT levels were slightly lower among current and former cigarette smokers than among nonsmokers.

Internal Mortality Comparisons

The distribution of all deaths and of cancer deaths according to tercile of serum DDT is displayed in Table 1. The slight positive trend in overall mortality with increasing serum DDT is not significant either with the ordinal scale or the continuous logarithm measurement. The adjusted RRs are slightly lower than the unadjusted estimates, principally because of confounding by age and gender. The adjusted mean difference of the logarithm of serum DDT for decedents and the living is 0.025 (95 per cent confidence intervals = −0.02, 0.07).

There is no consistent positive trend in cancer mortality according to serum DDT. The adjusted mean of the logarithm of serum DDT is very slightly lower among persons who died of cancer than it is among those who did not (mean difference = −0.01, 95 per cent CI = −0.08, 0.07).

There is a slight positive trend in respiratory cancer rate according to tercile of serum DDT. However, the point estimates were unstable and the trend test p-values were high. Furthermore, the adjusted mean difference of the logarithm of serum DDT for those who had died of respiratory cancer compared with those who had not is 0.01 (95 per cent CI = −0.12, 0.14). As anticipated, cigarette smoking is a strong determinant of respiratory cancer mortality. The RRs of respiratory cancer for ex-smokers and current smokers compared with non-smokers are 8.3 (0.9, 77) and 24 (3.1, 193), respectively.

TABLE 1—Number of Deaths, Cancer Deaths, and Respiratory Cancer Deaths and the Relative Mortality Rates According to Terciles of Serum DDT

DDT Tercile	Number of Subjects	Person-Years	Deaths	Mortality Rate[a]	Relative Rate[b]
1 (0–31 ppb)	305	2,780	57	2,050	1.0
2 (31–52 ppb)	308	2,722	72	2,645	1.2 (0.8, 1.7)[c]
3 (>52 ppb)	306	2,646	80	3,023	1.2 (0.9, 1.8)
Trend test—ordinal					p = .27
Trend test—continuous					p = .21

DDT Tercile	Cancer Deaths	Mortality Rate[a]	Relative Rate[b]
1	15	540	1.0
2	23	845	1.5 (0.8, 3.0)
3	16	605	0.9 (0.4, 2.0)
Trend test—ordinal			p = .80
Trend test—continuous			p = .96

DDT Tercile	Respiratory Cancer Deaths	Mortality Rate[a]	Relative Rate[b]
1	5	180	1.0
2	7	257	1.5 (0.5, 4.9)
3	7	265	1.8 (0.5, 6.2)
Trend test—ordinal			p = 0.34
Trend test—continuous			p = 0.77

[a]Per 100,000 person-years.
[b]All relative rates are adjusted for age, gender, race, years of schooling, and cigarette habit.
[c]95% confidence intervals shown in parentheses.

TABLE 2—The Number of Observed Deaths, Cancer Deaths, and Respiratory Cancer Deaths and the Standardized Mortality Ratios

Referent Experience	All Deaths			Cancer Deaths			Respiratory Cancer Deaths		
	Observed	SMR	95% CI	Observed	SMR	95% CI	Observed	SMR	95% CI
United States	209	86	(75, 98)	58	98	(74, 126)	21	122	(76, 187)
South Carolina	209	81	(70, 93)	58	104	(79, 134)	21	124	(77, 189)

Seventeen persons died from cancer of the gastrointestinal system. There is a slight inverse relation between gastrointestinal cancer mortality and the logarithm of serum DDT. There were 18 other cancer deaths. No specific site included a number of deaths sufficient to permit a separate analysis. The RRs of all these other cancers for the second and third terciles compared with the first are 1.0 (0.3, 3.1) and 0.5 (0.1, 2.0), respectively.

External Mortality Analyses

The number of deaths and the SMRs for all causes, for cancer, and for respiratory cancer are displayed in Table 2. The overall mortality experience of study subjects was significantly lower than that of comparable individuals in the general population. The overall cancer mortality rate of study subjects was similar to that of the population of the United States and of South Carolina. On the other hand, the SMR for respiratory cancer mortality rate is higher, but not statistically significantly so, among study subjects.

Discussion

This study has a number of strengths for the purpose of evaluating the relation between DDT and cancer mortality, and at least one limitation—its small size. Strengths are that the follow-up was nearly complete and that the ascertainment and classification of cause of death was done without knowledge of the subjects' DDT levels. Furthermore, serum DDT was measured at the beginning of the follow-up period and so is not subject to distortion because of the presence of disease. Thus, the ascertainment of the outcome and the classification of the exposure are free of bias.

The exposure to DDT of this study population began in the late 1940s when it was used for mosquito control, in household pesticides, and for control of agricultural pests. Body burdens of DDT probably resulted from absorption through the skin and lungs and by ingestion. Subjects probably had been exposed chronically to DDT for at least 25 years by 1975, so that a reasonably long induction period would have elapsed.

Our subjects were not selected because of high DDT exposures, but their serum DDT levels were generally higher than among other comparable persons in the southern region. That is, if the age-gender-race structure of the study subjects is applied to the corresponding age-gender-race specific mean values of the logarithm of total serum DDT among southerners in the Second National Health and Nutrition Examination Survey,[11] an expected mean of 1.46 (about 35 ppb) is obtained, whereas the mean among our subjects is 1.59 (about 48 ppb). As another comparison, among 37 decedents aged 55 and over who had been occupationally exposed to pesticides,[17] the mean level of total serum DDT was about 92 ppb whereas among subjects in our upper tercile the mean is about 75 ppb. About 6 per cent of our study subjects had serum DDT levels exceeding 92 ppb. Thus, a number of our subjects apparently had sustained high DDT exposures.

A limitation of the study is that serum DDT and DDE serve as a surrogate measure of past DDT exposures. To the extent that current serum DDT levels are not an accurate indication of past DDT exposures, there will be random misclassification of the exposure and hence a bias in the direction of an absence of association between DDT and cancer. This is of special concern in a negative study such as this, but since DDT is metabolized very slowly, the bias probably is small.[15,16]

The findings of this study do not support the hypothesis that exposure to DDT increases cancer mortality. There is no consistent dose-response between cancer mortality rates and serum DDT levels and the mean level of serum DDT is lower among subjects who had died of cancer than it is among those who had not. Based upon the PH model of cancer mortality with the logarithm of serum DDT as the exposure measure, an upper 95 per cent confidence limit for the increase in cancer mortality sustained by a doubling of serum DDT is about 35 per cent. Thus, although the study includes only 54 cancer deaths, its findings effectively rule out a 50 per cent or more increase in overall cancer mortality attributable to a two-fold increase in serum DDT levels.

The findings with respect to respiratory cancer are less clear. Although there is some evidence of a dose-response relation between serum DDT and respiratory cancer, the point estimates are unstable and the trend is not statistically significant. Furthermore, respiratory cancer risk is lower among subjects in the upper quintile than it is among persons in the middle tercile. The finding that the mean levels of serum DDT are nearly identical for persons who had and had not died of respiratory cancer also does not support a causal interpretation of the finding. Based upon the respiratory cancer PH model it is estimated that a doubling of serum DDT is associated with only an 8 per cent increase in respiratory cancer mortality. On the other hand, the study is too small to exclude moderate increases in respiratory cancer mortality attributable to DDT. The upper 95 per cent confidence limit for excess respiratory cancer mortality associated with a doubling of serum DDT is 83 per cent.

The findings regarding other cancers also suffer from the imprecision resulting from a small study size. Nonetheless, the finding of an inverse relation between cancers of the gastrointestinal tract and cancers from all other sites combined, as well as the observation that no cluster of any particular form of cancer is evident among persons in the upper tercile of serum DDT, provides some assurance that DDT is not related to human cancer. The external SMR analysis indicates that the overall cancer mortality rate in the cohort is about the same as it is in the general population whereas the lung cancer mortality rate may be slightly higher.

Other epidemiologic studies generally do not support the opinion that DDT causes human cancer. In a retrospective follow-up study (RFUS) of pesticide applicators, the overall cancer SMR was 72 with 95 per cent confidence limits of 53 to 96.[18] In the same study the lung cancer SMR was 115 (77,

170), but the risk of lung cancer death was unrelated to duration of employment or the intensity of pesticide exposures. In another RFUS of pesticide applicators,[5] the lung cancer SMR was 135 (94, 189) and there was a consistent dose-response relation between length of licensure as a pesticide applicator and the lung cancer SMRs. In an East German study of agricultural applicators, a two-fold excess of lung cancer was reported.[6] Limitations of these studies are: 1) that these men had been exposed to other chemicals and pesticides so that it is impossible to attribute an excess of lung cancer among them specifically to DDT, and 2) the lack of information on their smoking habits also detracts from a causal interpretation of the findings.

Studies pertaining specifically to DDT do not support the opinion that it is related either to all cancer or to lung cancer mortality. In a follow-up study of workers occupationally exposed to pesticides, there was little difference between the serum DDT levels of persons who developed cancer or respiratory tract cancer as compared to those who did not.[17] In a study of 354 men employed at a plant that manufactured DDT,[19] the SMR for all cancer was 68 (25, 147), while the SMR for respiratory cancer was 125 (34, 321). In another study of workers engaged in the manufacture of DDT,[20] the SMR for all cancer was 99 (82, 119), while the SMR for lung cancer was 149 (68, 282), but in a nested case-control study among these and other workers, the RR of respiratory cancer among men exposed to DDT compared with those not so exposed was 0.74 (95% CI = 0.2 to 2.4, our estimate), while the RR of respiratory cancer pertaining to exposure to inorganic bromides was 7.0 (95% CI: 1.5 to 33, our estimate). This observation suggests that the excess of respiratory cancer observed among these DDT workers may have been due to some exposure other than DDT.

Higginson recently reviewed the epidemiologic evidence pertaining to DDT and cancer and concluded that DDT has had no significant impact on human cancer.[21] The present study supports this opinion. Nonetheless our findings with respect to respiratory cancer are not reassuring. On the one hand, there is a weak and not entirely consistent dose-response relation between serum DDT levels and respiratory cancer mortality that may be due to chance. On the other hand, previous epidemiologic studies have implicated pesticides (but not specifically DDT) as a cause of human lung cancer. Unfortunately, the issue of whether or not DDT is a weak lung carcinogen, or is associated with an increased risk of cancer of some rare site, can be resolved only in a study much larger than the present one.

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A Birth Cohort Analysis of Dental Contact among Elderly Americans

FREDRIC D. WOLINSKY, PHD, AND CONNIE LEA ARNOLD

Abstract: We applied standard cohort and multiple regression techniques to data on the dental utilization rates of 129,191 elderly individuals taken from the 1972, 1973, 1976, 1977, 1980, and 1981 Health Interview Surveys. The results indicate that the marked variation in dental contact rates is a reflection of *cohort succession*, and not a function of aging *per se*. Older cohorts having lower dental contact rates are being replaced by younger cohorts having higher dental contact rates. The dental contact rates of the individual birth cohorts themselves are quite stable over time. The results also indicate that economic barriers (especially liquid assets) have become more important than ever before, especially for the oldest-old. These findings have important implications for public policy about the oral health and health care of elderly Americans. (*Am J Public Health* 1989; 79:47–51.)

Introduction

Dental care for elderly Americans is a significant issue that carries with it a complex web of financial, political, and ethical considerations and implications. Although numerous studies have been published on the epidemiology of oral health problems and the use of dental services among the elderly,[1–11] that literature has been methodologically limited by its reliance on single cross-sectional surveys. The consistently observed age-group differences reflect the effects of cohort differences *and* the aging process, which are inseparable in single, cross-sectional designs.[12–23] Thus, whether the consistently reported age-group decline in dental contact rates indicates aging, cohort, or aging and cohort effects remains unknown. Also unknown are plausible period (i.e., historical) effects on dental services utilization associated with the increasing relative supply of dentists and use of dental hygienists and assistants over the past several decades.[24]

The purpose of this paper is to apply the standard cohort and multiple regression techniques used elsewhere[25–27] to sequential, cross-sectional data pooled from the Health Interview Surveys (HIS) conducted by the National Center for Health Statistics (NCHS).[28,29] This will allow for the separation and examination of the effects of age, period, and cohort on the use of dental services by the elderly.

Methods

Data on 129,191 elderly individuals were taken from the 1972, 1973, 1976, 1977, 1980, and 1981 HISs conducted by the NCHS.[28,29] The HIS is an ongoing annual survey that follows a multistage probability design sampling the noninstitutionalized civilian population of the United States. Eight four-year birth cohorts were identified ranging from 56–59 years of age in 1972 to 92–95 years of age in 1980. These cohorts provide a legitimate opportunity for examining dental contact among the elderly.

Table 1 shows the age composition, cohort designation, per cent having seen a dentist in the past 12 months, and number of cases per cohort at each of the three points in time. Except for compositional changes within aging cohorts and the effects of sampling error, variation in the per cent having seen a dentist can be identified straightforwardly. Age effects may be determined by examining *intra*cohort differences (i.e., by reading diagonally down and to the right). Cohort membership effects may be determined by examining *inter*-cohort changes (i.e., by reading down the columns). Period effects may be determined by comparing the same age group at one point in time with that in another point in time (i.e., by reading across the rows).

There is, however, a confounding problem. Age and cohort effects are both represented in column comparisons, cohort and period effects are both evidenced in row comparisons, and age and period effects are both involved in diagonal comparisons. Thus, the separation of age, period, and cohort effects by visual inspection of standard cohort tables is difficult unless the observed effects are both pronounced and consistent across all comparisons. When this is not the case, more sophisticated statistical techniques must be used to resolve the identification problem.

As informative as the standard cohort analysis of the

TABLE 1—Age Composition, Cohort Designation (in parentheses), Number of Cases per Cohort per Year, and Per Cent Having Seen a Dentist in the Past 12 Months

Age Composition (years)	Survey Years		
	1972–73	1976–77	1980–81
56–59	9,839 (1)	8,910 (9)*	8,684 (9)
%	52.2	57.1	60.7
60–63	8,739 (2)	8,056 (1)	7,883 (9)
%	46.7	51.2	55.8
64–67	7,560 (3)	6,999 (2)	6,871 (1)
%	42.2	47.5	49.7
68–71	6,158 (4)	5,843 (3)	5,680 (2)
%	37.6	41.2	45.1
72–75	4,757 (5)	4,492 (4)	4,337 (3)
%	32.6	37.3	41.8
76–79	3,479 (6)	3,219 (5)	3,030 (4)
%	27.9	31.7	37.0
80–83	2,137 (7)	2,100 (6)	1,986 (5)
%	22.7	29.3	33.1
84–87	1,030 (8)	1,049 (7)	1,117 (6)
%	20.3	22.6	24.3
88–91	430 (9)	416 (8)	425 (7)
%	16.0	23.6	20.5
92–95	128 (9)	121 (9)	126 (8)
%	10.9	19.0	15.1
Totals	45,495	42,395	41,301
%	41.2	45.6	49.2

*For convenience, any cohort that enters or leaves the standard table is designated as cohort 9.

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data in Table 1 will prove to be, it does not take into consideration the effects of other factors known to affect dental services consumption. Therefore, a second analytic approach is also used. It involves using ordinary least squares in a multiple regression analysis[30,31] to estimate a predictive model of dental contact separately within each column (i.e., time period) of Table 1, and then comparing the observed effects across columns.

The multivariate approach selected is based on Andersen's behavioral model of health services utilization,[32] which has been frequently applied to measures of dental contact.[3,6,11] It defines the use of health services as a function of the predisposing (age, marital status, living arrangements, race, education, and labor force participation), enabling (family income, region of residence, and population density), and need (self-assessed health status and the extent of activity limitations due to health conditions) characteristics of the individual.

The Appendix shows (for all cohorts during 1980–81) the means, standard deviations, and coding algorithms of the indicators used to measure the predisposing, enabling, and need characteristics. The need measures reflect individuals' global health status assessments, and thus are not specific to their need for dental care. Although direct indicators of dental need would be preferable, they are not available for the HISs. Prior multivariate assessments have found specific indicators of dental problems to be statistically unrelated to dental contact[6]; therefore, these indicators would seem appropriate for the present study.

Results

The data in Table 1 indicate three distinct patterns. First, column comparisons reveal consistent, monotonic age-group decreases within each of the pooled surveys. When coupled with the increases in the grand (i.e., within column) means (a short-hand method of partitioning out the confounding effects of aging, which also affect column comparisons), these data suggest a rather strong pattern of change resulting from cohort succession. That is, as new cohorts enter and older cohorts exit the table, the overall percentage of individuals aged 56 to 96 years old who have been to the dentist in the past year increases from 41.2 per cent in 1972–73, to 45.6 per cent in 1976–77, and to 49.2 per cent in 1980–81. Thus, in less than a decade, cohort succession has raised the grand mean of dental contact by a full eight percentage points.

The second pattern involves row comparisons. With two exceptions a consistent monotonic increase is observed across the rows. The exceptions involve the 88–91 and 92–95 year old age groups in 1976–77 and 1980–81. For these age groups about a three or four percentage point drop in dental contact can be observed between columns 2 and 3. In general, however, the row comparisons appear to suggest a period effect that results in higher rates of dental contact over time. Although this interpretation would be consistent with the increasing relative supply of dentists during this same period,[24] it must be remembered that row comparisons also reflect cohort succession, and the column comparisons indicated a clear and consistent pattern of cohort effects. Thus, the row comparisons most likely are artifacts of the effects of cohort succession.

The diagonal comparisons provide the third pattern, although it is not as straightforward. For cohorts 2, 3, 4, and 5 the diagonal (i.e., aging) comparisons indicate relatively stable dental contact rates. Only cohort 1 exhibits a consistent decline, with its contact rate falling by 3.5 percentage points between 1972–73 and 1980–81. The three oldest cohorts (6, 7, and 8) experience either a stable or modest increase between 1972–73 and 1976–77, followed by a larger decline between 1976–77 and 1980–81. Over the full eight-year period, the dental contact rates among these three oldest cohorts decline by between 2.2 and 5.2 percentage points. Thus, the most consistent pattern among the diagonal comparisons is the absence of any aging effects. This provides further support for the dominance of cohort succession effects as the principal explanation of the variation in dental contact rates shown in Table 1.

Table 2 contains the results of the regression analysis. Here age is measured by a set of dummy variables representing each of the cohorts through time. Cohort 1 is the reference category, with its dummy variable omitted from the regression analysis. The interpretation of the cohort dummies within columns of Table 2 is equivalent to the cross-sectional cohort assessment made within column comparisons in Table 1.

The interpretation of the cohort dummies across columns (within rows) in Table 2, however, represents the change in the effect of belonging to a particular cohort as that cohort ages. This is a more pristine distillation of *aging* effects than that shown in the diagonal comparisons of Table 1 for two reasons. First, the aging effects shown here are *net* effects, with the joint effects of the predisposing, enabling, and need characteristics having already been partitioned out. Second, the multiple regression design is such that some of the associated period effects have been residualized to the intercept. Therefore, any differences observed by comparing, for example, the partial, unstandardized regression coefficient of cohort 2 in column 1 with that shown for it in column 2 more closely reflect the effect of cohort 2's having aged four more years. Thus, if the decline in dental contact across age groups reflects the effects of the aging process, there should be a monotonic increase in the absolute value of the coefficients within the rows for each of the cohort dummies. In contrast, if the decline in dental contact across age groups reflects cohort differences, then the coefficients for the cohort dummies should be relatively stable within the rows.

The data in Table 2 clearly show that the partial, unstandardized regression coefficients for the cohort dummies do not exhibit a consistent, monotonic pattern of change. Indeed, the coefficients for cohorts 2, 3, 4, and 5 are virtually isomorphic within rows. Where they do fluctuate, these changes are marginal (i.e., within rows, each coefficient falls within each other coefficient's 95 per cent confidence interval). Thus, at least for these cohorts, the traditionally observed decline across age groups appears to be the result of *cohort succession*, with older cohorts having less dental contact than their younger counterparts.

It is only for cohorts 6, 7, and 8 that significant changes in the coefficients are observed within the rows. This might be taken as an indication that dental contact rates decline with age, but only among the oldest cohorts. If this were the case, however, a consistent, monotonic pattern of change would be expected across both columns for the oldest cohorts. But all of the changes occur between columns 2 and 3, and only for cohorts 6, 7, and 8. Thus, the data indicate statistical interaction between age and period, resulting in a pattern of effects showing a consistent decline in dental contact rates among cohorts who have aged into their eighth decade or beyond between 1976–77 and 1980–81.

TABLE 2—The Unstandardized Regression Coefficients (and their standard errors) Obtained by Predicting the Annual Rate of Dental Contact, for all Cohorts Combined, per Year

Independent Variables	Survey Years		
	1972–73	1976–77	1980–81
Sex	.042 (.005)	.045 (.005)	.037 (.005)
Widowhood	−.061 (.007)	−.067 (.008)	−.056 (.008)
Living alone	.081 (.007)	.087 (.008)	.080 (.008)
Race	−.012 (.008)	−.036 (.009)	−.020 (.009)
Education	.025 (.001)	.027 (.001)	.028 (.001)
Labor force participation	.006 (.006)	.008 (.006)	.001 (.006)
Cohort 2	−.021 (.007)	−.005 (.008)	−.026 (.009)
Cohort 3	−.033 (.008)	−.047 (.009)	−.034 (.010)
Cohort 4	−.058 (.008)	−.060 (.010)	−.060 (.011)
Cohort 5	−.086 (.009)	−.093 (.011)	−.092 (.013)
Cohort 6	−.116 (.010)	−.112 (.013)	−.155 (.017)
Cohort 7	−.149 (.012)	−.161 (.017)	−.188 (.026)
Cohort 8	−.160 (.017)	−.146 (.027)	−.218 (.049)
Cohort 9	−.209 (.023)	−.008 (.007)	.021 (.007)
Family income	.078 (.003)	.100 (.003)	.141 (.005)
Northeastern US	.035 (.006)	.021 (.007)	.026 (.007)
Southern US	.032 (.006)	.037 (.006)	.017 (.007)
Western US	.057 (.007)	.061 (.008)	.045 (.008)
Major city residence	.038 (.005)	.046 (.005)	.033 (.005)
Farm residence	−.024 (.010)	−.012 (.012)	−.030 (.014)
Perceived health status	.033 (.006)	.041 (.006)	.040 (.006)
Limited activity	−.006 (.006)	−.015 (.005)	−.011 (.006)
Intercept	.030 (.011)	−.033 (.012)	−.024 (.013)
R^2	.14	.15	.15
$\bar{X}$ dental contact rate	.41	.46	.50

Discussion

These analyses have shown that the traditionally observed decline in dental contact rates across age groups is *not* a function of the aging process. Rather, it is a function of *cohort succession*. Generally speaking, dental contact rates are quite stable within birth cohorts as they age from their mid-50s to their mid-90s. Younger cohorts simply have higher rates of dental contact than older cohorts. As a result, the analysis of single, cross-sectional surveys has been somewhat misleading.

The identification of the effects of cohort succession suggests that if the trends observed in the 1970s continue through the 1980s, then by the early 1990s dental contact rates among those aged 56–96 should rise by about another 10 percentage points or so. That would result in three out of five older Americans going to see the dentist at least once a year. Thus, without any programmatic intervention, the oral health care of the elderly should continue to improve. The identification of the cohort succession effects also suggests that dental contact is a lifestyle trait that is relatively stable over the life course. As such, it would seem that oral health education programs targeted at older persons are unlikely to be effective in altering the behavior of those who routinely do *not* see a dentist on an annual basis.

The R^2 values of Table 2 (0.14 to 0.15) are generally consistent with prior studies of dental utilization among elderly Americans. Nevertheless, the low values underscore the fact that far too much oral health behavior remains unexplained. Better measurements, and the inclusion of additional indicators, such as those tapping dental insurance coverage, would seem fruitful for use in future research. Extant studies, however, offer little hope that such modifications will appreciably enhance the R^2 levels.[3,6,11,34]

With the exception of income, the effects of all of the other factors in the behavioral model (besides age) are quite stable over time. Being female, living alone, having more education, living outside of the midwest, living in a major metropolitan area, having more income, and being in good health all increase the likelihood of having seen a dentist in the past year. Being widowed, being a minority, and living on a farm all decrease the likelihood of dental contact. These results are also generally consistent with those of previous assessments.

It is the increasing importance of income over time that warrants further consideration. In 1972–73 the effect of income was such that each $10,000 increment increased the probability of dental contact by about 0.08. By 1980–81 that effect had nearly doubled, increasing the probability by about 0.14. These differences are substantively important. They indicate that economic barriers to dental care not only failed to dissipate, they actually increased considerably during the 1970s.

These cohort analyses have also shown isolated interaction effects. In general, their explanation is both difficult and requires heavy reliance on historical context.[12] These particular interaction effects (resulting in declining rates of dental contact) involved only those birth cohorts who had reached their eighth decade or beyond between 1976–77 and 1980–81. It was precisely during this period that double-digit inflation occurred. It likely affected the oldest-old more than their younger counterparts, because their general economic wealth is substantially lower,[35] and their assets are disproportionately less liquid.[36]

Given that the consumption of dental services is generally viewed as the most discretionary form of health services utilization,[37] it seems quite plausible that the oldest-old may

have been forced to cut back on dental care in order to maintain their consumption patterns of more essential goods and services. And that suggests that economic factors (especially in the form of liquid assets like income) can become (or may already be) critical barriers between elderly Americans and the dental care that they need. As a result, the oral health component of the American health care delivery system is inequitable.

ACKNOWLEDGMENTS

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APPENDIX

Coding algorithms, means and standard deviations for the indicators of the predisposing, enabling, and need characteristics in the 1980–81 Health Interview Surveys, all cohorts combined

Indicator	Algorithm	Mean	Standard Deviation
Predisposing Characteristics			
Sex	1 = female 0 = male	.56	.50
Widowed	1 = yes 0 = no	.24	.42
Lives alone	1 = yes 0 = no	.23	.42
Race	1 = nonwhite 0 = white	.09	.29
Education	number of years	10.4	3.65
Labor force participation	1 = yes 0 = no	.32	.47
Cohort 2	1 = yes 0 = no	.14	.35
Cohort 3	1 = yes 0 = no	.11	.31
Cohort 4	1 = yes 0 = no	.07	.26
Cohort 5	1 = yes 0 = no	.05	.21
Cohort 6	1 = yes 0 = no	.03	.16
Cohort 7	1 = yes 0 = no	.01	.10
Cohort 8	1 = yes 0 = no	.00	.05
Cohort 9	1 = yes 0 = no	.41	.49
Enabling Characteristics			
Family income	in tens of thousands of 1979 dollars	1.15	.67
Northeast region	1 = yes 0 = no	.24	.43
Southern region	1 = yes 0 = no	.32	.47
Western region	1 = yes 0 = no	.18	.38
Major city residence	1 = yes 0 = no	.37	.48
Farm residence	1 = yes 0 = no	.03	.18
*Need Characteristics**			
Perceived health status (good)	1 = yes 0 = no	.71	.45
Limited activity (some)	1 = yes 0 = no	.39	.49

*Following recent suggestions[33] both measures are dichotomized to indicate self-assessments of good vs less than good health (or no vs some activity limitations).

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New Technique Saves Blood Usually Lost During Surgery

Researchers at Northwestern Memorial Hospital in Chicago have developed an innovative procedure for retrieving increased amounts of blood usually lost during surgery, thereby reducing the need for transfusions of donor blood. The technique entails recycling blood saved from absorbent sponges used during surgery.

Most hospitals that recycle the patient's own blood (called "autologous") do so by cleansing and reinfusing blood suctioned from the operating field during surgery. In the past, blood-soaked surgical sponges were weighed to help the operating team compute the amount of blood lost during surgery and then thrown away.

Studies have found that suctioning saves almost 50 per cent of the blood usually lost during surgery. However, the Northwestern group—believing that the blood absorbed by the sponges also could be reinfused during surgery—developed a technique of soaking surgical sponges in saline and filtering, washing, and reinfusing red blood cells. The technique was successful in retrieving an additional 40 per cent of red blood cells that would have otherwise been lost.

Because of health concerns about transfusing donor blood, there is a growing interest in using autologous blood to replace that lost during surgery. Many patients give their own blood several weeks before their surgery to ensure a supply for transfusion.

The NMH surgeons first replace lost blood with that retrieved from suctioning and from the sponges before using blood donated by the patient, and finally units donated by others. The sponge and suctioning retrieval systems are used in procedures involving considerable blood loss, such as in orthopedic and cardiac surgery. In the presence of infection or malignancy, however, intraoperatively retrieved autologous blood cannot be used.

Changes in Alcohol-Related Inpatient Care in Stockholm County in Relation to Socioeconomic Status during a Period of Decline in Alcohol Consumption

ANDERS ROMELSJÖ, MD, PHD, AND FINN DIDERICHSEN, MD, PHD

Abstract: Alcohol sales in Stockholm County decreased by 18 per cent from 1976 to 1981. The socioeconomic status of inpatients treated for alcohol psychosis, alcoholism, alcohol intoxication, liver cirrhosis, and pancreatitis was studied by linking data from the National Housing and Population Censuses in 1975 and 1980 with the inpatient care registers for 1976 and 1981. In both years, all rates were highest for people outside the labor market and lowest among white collar employees. The employment rate for those aged 25–44 years and treated in 1981 for alcohol psychosis, alcoholism, and alcohol intoxication—already low in 1975—had drifted further downward by 1980. Total rates of inpatient treatment for alcohol-related diagnoses generally declined but the gap between blue collar workers and white collar workers widened. We conclude that the goal for national alcohol policy, suggested by the WHO—a reduction of per capita consumption—should be combined with additional measures that will reach all social groups. (*Am J Public Health* 1989; 79:52–56.)

Introduction

Alcohol sales in Sweden increased from 4.9 liters of 100 per cent ethanol per inhabitant aged 15 and over in 1954 to a peak of 7.7 liters in 1976, but then decreased to 6.3 liters in 1981.[1,2] Alcohol sales in Stockholm County have exceeded national levels, but have also declined 18 per cent since 1976. Patients hospitalized for alcohol-related medical conditions have declined,[1,2] as well as charges for drunkenness and drunken driving in Stockholm County.[3–5] However, there has been no decline in deaths (both sexes) or hospitalization of women due to alcohol psychosis, alcoholism, and alcohol intoxication up to 1984.[3]

Survey data indicate that since the late 1960s, alcohol consumption has increased most in social class III (mainly blue collar workers), possibly because the increased equalization of disposable income from 1967 to 1981 made alcoholic beverages relatively cheaper for blue collar workers.[6] Surveys since the mid-1970s find that frequent intoxication among military conscripts has been more common among those with a low formal education and that its decline from 1978–79 to 1983–84 has been greater among military conscripts with a higher formal education.[4]

We hypothesized that the difference in rates of alcohol-related inpatient care between the manual and the non-manual workers categories has increased. We therefore measured the socioeconomic distribution of alcohol-related inpatient care during a period of decline in alcohol sales. A second aim was to elucidate social mobility among inpatients treated for alcohol-related problems.

Methods

Study Population

Stockholm County is a metropolitan area with a population of around 1.5 million inhabitants in 1975 as well as in 1980. At least 97 per cent of all discharges from hospitals in Stockholm County have been registered in the Inpatient Care Register (ICR),[7] making the ICR valuable for epidemiological studies. The Maria Clinic, which accounts for about one-third of all discharges under the diagnosis of alcoholism in the county, is not included in the ICR. At least 95 per cent of the discharges at the Maria Clinic concern patients living in its geographically limited catchment area.

In Sweden, Population and Housing Censuses have taken place every fifth year. The censuses of 1975 (FoB75) and 1980 (FoB80) were compulsory,[8,9] with a non-response rate less than 1 per cent. The definitions of the variables used in this study from the two censuses were essentially identical.[9] A tape was produced by the National Board of Health and Welfare by linking data in FoB75 for Stockholm County with the ICR for 1976 with the aid of the personal identity number which every Swedish citizen has been assigned.[10] The same procedure was performed with FoB80 and the ICR for 1981.

Diagnoses and Social Variables

We studied alcohol psychosis, alcoholism, alcohol intoxication, and the closely alcohol-related somatic diseases of liver cirrhosis and pancreatitis,[11] pooling data for the first three diagnoses because they overlap,[12] and abbreviated them to "AAA" in this report. Within a specific diagnostic category an individual discharged more than once has been counted only once, but one individual could be an inpatient in each of the three diagnostic categories. The study has been restricted to individuals 15–64 years of age for "AAA" and to individuals 25–64 years of age for the latter two categories; each sex has been studied separately. We grouped occupations on the basis of the most detailed level available in both censuses,[13,14] reaching a 90 per cent agreement with the new socioeconomic classification in FoB80. The categories were: manual (blue-collar) workers; non-manual employees, lower level; non-manual employees, intermediate and higher level, including self-employed persons; disability pensioners; house-wives; and "others" (unemployed, employed persons who were on sick leave or other leave since January 1, 1975 or January 1, 1980; those undergoing vocational rehabilitation; those who had had an occasional job during the year; family workers 15 hours or less). Non-manual employees and self-employed were combined into one category for the calculation of rates. Students and persons employed part time were excluded because numbers were small.

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Analyses

We calculated the rates of inpatients in 1976 and in 1981 for the whole population and for each socioeconomic category, using a 5 per cent random sample of the population in Stockholm County for the denominators. The rates and 95 per cent confidence intervals were calculated for 1975 and 1980 after standardization to the age-distribution in 1975 for each sex.[15] Since data on inpatient care for "AAA" did not include the Maria Clinic, we excluded persons living in the catchment area of the Maria Clinic from the denominator for "AAA". We estimated social mobility by comparing socioeconomic category for 1975 and 1980 among the 1981 inpatients aged 25–44 years in 1975 who participated in the 1980 census. Since downward social mobility could be caused by disability and ill health,[16,17] we attempted to measure change from a higher to a lower socioeconomic category between 1975 and 1980. For non-manual employees at an intermediate and higher level, downward drift means a transition to lower level: non-manual work, manual work, disability pension or to the "other" category. Downward drift for a manual worker means a transition to disability pension or to "other".

Results

Persons outside the labor market had the highest rates for all of the alcohol-related diagnoses for both sexes and in both years (Table 1). Rates were higher for men than for women, and for manual workers than non-manual employees. Among male non-manual employees, the rates were higher for those at the lower social level (Figure 1). The changes from 1976 to 1981 varied according to diagnosis (Table 1). Among the employed the decrease was generally greater, or the increase smaller, among non-manual employees than among manual workers in both sexes (Figures 1–2). For both sexes, the rates increased with age (Table 2). The change from 1976 to 1981 also varied with age, with the greatest changes being in younger ages. There was an increase for "AAA" among women.

Figures 3 and 4 show a considerable downward drift for in-patients with "AAA" in all socioeconomic categories. Thirty-four per cent of the male inpatients who were employed in 1975 had become disability pensioners or left the labor market by 1981. The corresponding figure for women was 29 per cent. The percentage becoming disability pensioners or "others" did not generally depend on the patient's 1975 socioeconomic category. (Housewives, students and part-time employees are excluded from the comparisons.)

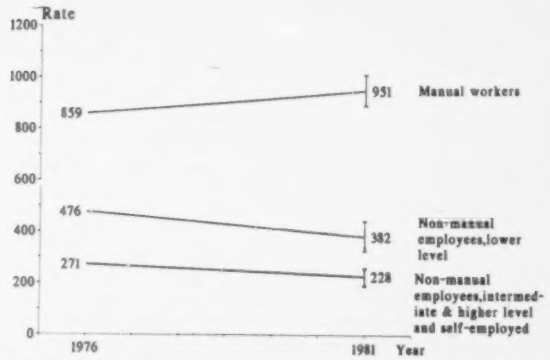


FIGURE 1—Male Inpatients Ages 25–64 Years Treated for Alcohol Psychosis, Alcoholism and Alcohol Intoxication in 1976 and 1981 among Manual Workers and Non-manual Employees (Age-standardized rates per 100,000 men and 95% confidence intervals for standardized rate ratios)

Discussion

Our estimates may reflect biases in the use of diagnostic criteria, selection to inpatient care, criteria for disability pensioning, health-related social mobility, or factors related to working conditions and alcohol availability.

The Use of Diagnostic Criteria—We issued a short questionnaire to the head physicians of all departments of internal medicine, surgery, orthopedics, psychiatry, and

TABLE 1—Age- and Sex-standardized Rates per 100,000 Inhabitants for In Patients Ages 25–64 Years with Alcohol Psychosis/Alcoholism/Alcohol Intoxication ("AAA"), Liver Cirrhosis and Pancreatitis in 1976 Distributed by Socioeconomic Category in 1975 (Standardized rate ratios (SRRs)—rates in 1981 in relation to rates in 1976—with 95% confidence intervals)

	"AAA"			Liver Cirrhosis			Pancreatitis		
	Rate 1976	SRR (1981/1976)	95% CI	Rate 1976	SRR (1981/1976)	95% CI	Rate 1976	SRR (1981/1976)	95% CI
Men									
n 1976	(2,851)			(248)			(476)		
Manual workers	859	1.11	(1.04, 1.18)	31	0.88	(0.64, 1.22)	110	0.78	(0.65, 0.93)
Non-manual employees and self-employed	351	0.79	(0.71, 0.88)	25	0.71	(0.51, 0.99)	63	0.67	(0.54, 0.83)
Disability pensioners	4,642	0.87	(0.77, 0.98)	392	0.54	(0.37, 0.80)	303	1.35	(0.95, 1.91)
"Others"	4,577	0.98	(0.92, 1.05)	292	0.74	(0.62, 0.83)	521	0.74	(0.60, 0.89)
Average	1,055	0.98	(0.94, 1.02)	65	0.71	(0.62, 0.83)	127	0.75	(0.68, 0.83)
Women									
n 1976	(652)			(133)			(147)		
Manual workers	182	1.39	(1.21, 1.61)	20	0.90	(0.58, 1.41)	23	0.65	(0.46, 0.92)
Non-manual employees and self-employed	136	1.09	(0.94, 1.27)	10	1.73	(1.17, 2.56)	28	0.50	(0.33, 0.74)
Housewives	188	1.01	(0.74, 1.36)	27	0.61	(0.25, 1.48)	39	0.28	(0.10, 0.78)
Disability pensioners	923	1.23	(0.94, 1.61)	221	0.57	(0.37, 0.89)	81	0.53	(0.29, 0.97)
"Others"	1,419	1.06	(0.92, 1.22)	181	0.62	(0.37, 1.02)	145	0.69	(0.42, 1.12)
Average	247	1.18	(1.10, 1.26)	43	0.65	(0.54, 0.80)	38	0.63	(0.51, 0.77)

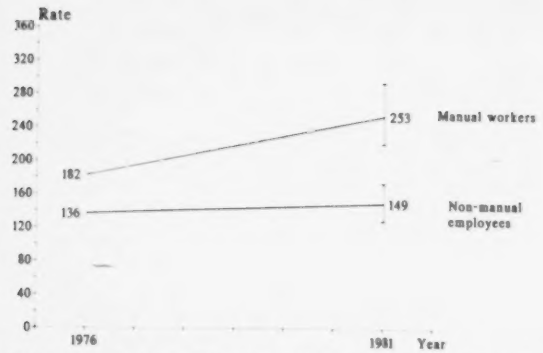
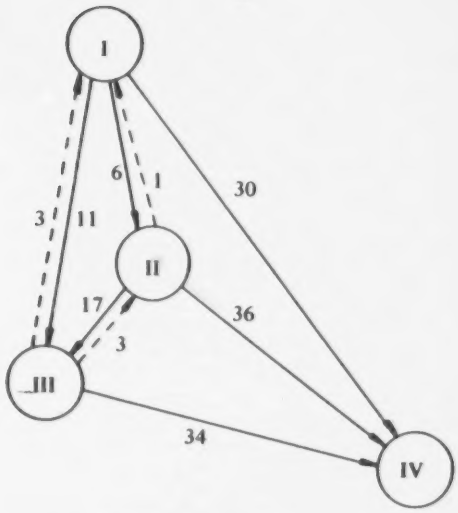


FIGURE 2—Female Inpatients Ages 25–64 Years Treated for Alcohol Psychosis, Alcoholism and Alcohol Intoxication in 1976 and 1981 among Manual Workers and Non-manual Employees (Age-standardized rates per 100,000 women and 95% confidence intervals for standardized rate ratios)

alcoholic diseases in the county, asking for any changes in criteria during 1973–83 for the diagnoses under study. According to the answers from the responding 75 per cent, there had been no such changes. There may have been some social bias in the use of the diagnosis "alcoholism" (and the related diagnoses "alcohol psychosis" and "alcohol intoxication"). Common screening tests (not chemical tests) in the United States have been considered to be biased towards the lower social classes,[18] but the similar patterns we observed for "AAA", liver cirrhosis and pancreatitis suggest that such factors have not been important in this study.

Outpatient Treatment—The changes in inpatient treatment between 1976 and 1981 may have been affected by the increase in outpatient visits from 132,000 to 141,000 visits and greater use of sensitizing agents. These factors were discussed in more detail elsewhere.[3] We lack data on the relation to sex, age, and socioeconomic status for most of those factors, but the rate differences between manual workers and non-manual employees could reflect, in part, greater use of outpatient care and treatment facilities by non-manual employees.

Stefansson and Cullberg have reported that a substantial expansion of psychiatric outpatient treatment facilities in two municipalities in Stockholm County in the mid-1970s was followed by a substantial increase in the proportion of the population receiving psychiatric outpatient care including patients with alcoholism.[19] The total consultation rate, related to population size, was greater for patients from social class III, but greatest for social class II. There was a decrease in hospital admissions for all social classes. Data on social class distribution for outpatients and inpatients with alcoholism were not presented. The proportion of all outpatient visits to doctors was roughly the same for all social classes in 1974 and in 1981 in Sweden.[20] Thus, we have no indication of changes in the social pattern of outpatient visits for alcohol-related problems in Stockholm County from 1976 to 1981, but the data are scanty.



I Non-manual employees, intermediate & higher level and self-employed, n = 149 in 1975

II Non-manual employees, lower level, n = 128 in 1975

III Workers, n = 680 in 1975

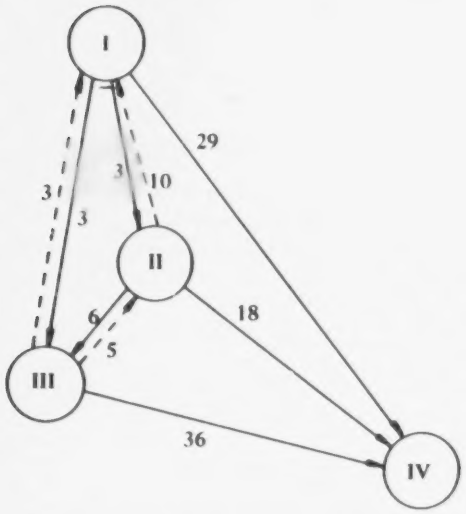
IV Disability pensioners and »others»

FIGURE 3—Social Mobility from 1975 to 1980 for Male Inpatients Ages 25–44 Years in 1981 with Alcohol Psychosis, Alcoholism and Alcohol Intoxication, Who Were Gainfully Employed in 1975

Disability Pensions—After new legislation in the beginning of the 1970s it became easier to get a disability pension, and from 1977 on to get disability pension due to alcoholism as a main diagnosis.[21] Alcohol-related diagnoses were thus

TABLE 2—Age- and Sex-specific Rates per 100,000 Inhabitants for Inpatients with Alcohol Psychosis/Alcoholism/Alcohol Intoxication ("AAA") Liver Cirrhosis and Pancreatitis in 1976 Distributed by Age in 1975 (Standardized rate ratios (SRRs)—rates in 1976 in relation to rates in 1981—with 95% confidence intervals)

	"AAA"			Liver Cirrhosis			Pancreatitis		
	Rate 1976	RR (1981/1976)	95% CI	Rate 1976	RR (1981/1976)	95% CI	Rate 1976	RR (1981/1976)	95% CI
Men									
15–24 years	257	0.41	(0.32, 0.52)						
25–44 years	933	0.90	(0.85, 0.94)	31	0.57	(0.42, 0.77)	111	0.67	(0.58, 0.78)
45–64 years	1,216	1.04	(0.98, 1.10)	113	0.77	(0.65, 0.91)	148	0.76	(0.66, 0.89)
Women									
15–24 years	106	0.80	(0.57, 1.10)						
25–44 years	245	1.14	(1.04, 1.25)	16	0.69	(0.46, 1.03)	31	0.58	(0.43, 0.79)
45–64 years	251	1.22	(1.10, 1.37)	58	0.76	(0.60, 0.96)	48	0.67	(0.51, 0.88)



I Non-manual employees, intermediate & higher level and self-employed, n = 34 in 1975

II Non-manual employees, lower level, n = 82 in 1975

III Workers, n = 176 in 1975

IV Disability pensioners and »others»

FIGURE 4—Social Mobility from 1975 to 1980 for Female Inpatients Ages 25–44 Years in 1981 with Alcohol Psychosis, Alcoholism and Alcohol Intoxication, Who Were Gainfully Employed in 1975

artefactually increased among disability pensioners, but the extent of this effect is not clear.

The total number of newly granted temporary or permanent disability pensions in Stockholm County decreased from 9,071 during the period 1973–75 to 7,881 during the period 1978–80, partly because the retirement age was lowered from 67 to 65 years in 1976.[22] We lack data on disability pensioning for alcoholism in Stockholm County, but we have data for Sweden as a whole. The number of newly granted disability pensions in Sweden with alcoholism or alcohol psychosis as the main diagnosis was 18 per cent higher in 1978–80 than in 1974–76 among men and 25 per cent higher among women (Information from Lena Eriksson, the National Social Insurance Board). Less than 5 per cent of all newly granted disability pensions in men and less than 1 per cent in women were due to alcoholism or alcohol psychosis as the main diagnosis in Sweden, during the 1970s. However, many alcoholics receive a disability pension under other main diagnoses.[23,24] The decline in disability pensioning and the small number of disability pensioners with a diagnosis of alcoholism suggest that changes in legislation and routines do not markedly bias a comparison of socioeconomic categories in 1976 with those in 1981.

Health-related Social Mobility—The much higher rates of alcohol-related inpatient care for disability pensioners and "others" can be explained in part by downward social drift of people unable to work due to alcohol-related disabilities.[16,17] Perhaps the educational requirements and experience needed to reach some of the positions in the category of non-manual work also may be more incompatible with alcohol abuse. In a 40-year follow-up of the Core City sample in Boston, Vaillant found a correlation between progressive course for alcohol abusers and membership in social class IV or V and concluded "that low social class does not cause alcoholism as much as severe alcoholism causes low social class."[25]

On the other hand, two Swedish studies suggest the opposite. Öjesjö, Hagnell and Lanke[26] found that the cumulative probability of developing alcoholism for males in the Lundby cohort in Sweden from 1957 to 1972 was roughly twice as high for manual workers as for middle strata. (Data on social class concerned 1957.) Romelsjö and Ågren[27] found that the mortality in Sweden from alcohol psychosis, alcoholism, alcohol intoxication, and liver cirrhosis was higher among manual workers than among other main occupational categories, but highest among disability pensioners and those who had left the labor market. Poikolainen[28] has reported that the relative risk for inpatient care among men aged 25–64 years in Finland during 1970–75 treated for alcohol psychosis, alcoholism, and alcohol intoxication was higher for manual workers than for clerical workers while the relative risk for liver cirrhosis and pancreatitis showed smaller differences between occupational groups.

Working Conditions—The differences in alcoholism among occupations may also reflect differences in the working environment, possibly because occupations with a greater risk of stress and monotony also carry a greater risk of alcohol abuse. Alfredsson, Spetz and Theorell[29] reported such correlations in the same linked data from 1975–76 used in the present study. Other studies have shown an increased risk of alcohol-related problems in specific occupations,[30,31] but it is not clear whether socioeconomic status relates to alcohol availability.

In sum, available data do not indicate any significant bias in the comparison of rates for various categories and for comparisons over time, and differentials between blue collar and white collar workers in alcohol-related disease appear to have widened. Reduced alcohol consumption has been advocated by the World Health Organization[32] and many countries—among them Sweden[33]—as an important goal for national alcohol policy. Our data show that although the decrease in sales and consumption was followed by a general decrease in alcohol-related problems,[1–4] the decreases were far from uniform across social groups. It seems necessary to combine a further reduction in alcohol consumption with additional measures to reach all social groups.[32–34]

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Publication Announcement: Child Studies Survey Available

A Survey of Sources for the Study of the History of Child Studies at the Rockefeller Archive Center (86 pp.) is the latest publication in a series of subject surveys of manuscript sources located at the Rockefeller Archive Center. Compiled by archivist Melissa A. Smith, the survey is available free of charge to interested researchers.

Over the past 80 years, American philanthropic foundations have responded to issues ranging from clean milk campaigns to jobs for unemployed youth. These two diverse causes illustrate the scope of these philanthropies' concern for the health, safety, mental and social development of children in American society.

Materials from the Commonwealth Fund, the General Education Board, and the Laura Spelman Rockefeller Memorial collections comprise the bulk of the survey. Other collections represented include those of the Rockefeller Brothers Fund, the Rockefeller Family, the Russell Sage Foundation, and the Bureau of Social Hygiene. The survey lists folder titles, box and folder numbers, and, when available, brief descriptions of the folder contents.

Researchers and institutions who want a copy of the free survey should contact the Director, Rockefeller Archive Center, 15 Dayton Avenue, Pocantico Hills, North Tarrytown, NY 10591-1598; Tel: (914) 631-4505.

A Study of Battered Women Presenting in an Emergency Department

Susan V. McLeer, MD, and Rebecca Anwar, PhD

Abstract: Medical records of female trauma patients were retrospectively reviewed prior to introducing a protocol for enhancing identification of battered women; 5.6 per cent of 359 female trauma patients were identified as having injuries caused by battering; 30 per cent of 412 patients following use of the protocol were so identified. Standardized protocols for identifying battered women among female trauma patients should be instituted in emergency departments. (*Am J Public Health* 1989; 79:65–66.)

Introduction

Awareness of the extent of domestic violence has heightened since the first reports appeared in the literature during the early 1970s.[1–4] Multi-disciplinary systems have evolved to protect victims, reduce violence, and expand resources.[5] Over 700 shelters have been established, and public policies have evolved at local, state, and federal levels. Nonetheless, in a national survey of 2,143 homes, battering was reported by more than 25 per cent of couples.[6] In a Harris poll of 1,793 women in Kentucky, 21 per cent reported being physically abused, with two-thirds of those divorced or separated being battered.[7] Similar findings were reported in Pennsylvania (26 per cent) and Texas (29 per cent).[8,9] Battering appears to be a more frequent event than job-related assaults of police officers.[10,11]

The National Crime Survey (NCS) projected that in 1980 there were 192,000 assaults within families. While these figures of reported crimes are considerably lower than those determined by other studies (estimating 1.8 million[6] to 3–4 million[12] battered women/year), the NCS data, nonetheless, provide a sobering analysis of the annual morbidity associated with domestic violence: 21,000 hospitalizations, 99,800 days of hospitalization, 28,700 emergency department visits, and 39,900 visits to physicians. The total annual health care costs approximate $44,393,700 with indirect costs of 175,500 lost days from paid work.[13]

These estimates do not begin to address the pain and misery associated with chronic victimization, nor the emotional sequelae for children who witness this form of violence. The mortality is significant also with 17 per cent of the nation's homicides occurring within the family, with half of these deaths being secondary to spouse killing a spouse.[14]

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Stark and Flitcraft reported that abuse may be the most common cause of injury in women,[15] with more injury episodes than auto accidents, muggings, and rape combined. Flitcraft in an early study of battering developed the following protocol for identifying and classifying battered women.[16]

	Classification of Battered Women
Positive:	**Injury was attributed to spouse or boyfriend in the medical record of the event.**
Probable:	**The record reports that the patient was beaten, kicked, hit, punched, but no personal etiology was noted.**
Suggestive:	**The recorded etiology of the injury did not seem to account adequately for the injury.**
Negative:	**Nothing in report of the injury would raise suspicion that injury was a result of battering.**

Using this classification system, we reviewed the records of all female trauma exclusive of motor vehicle accidents and natural disasters registered in the Emergency Department (ED) of the Medical College of Pennsylvania (MCP) during 1976. Among the women's injuries, 5.6 per cent were classified as positively caused by battering; 10.9 per cent were classified as probable; and 9.2 per cent as suggestive. Since previous reports had indicated that battered women were often times secretive and reluctant to report when injuries were caused by their having been beaten,[2] we designed a pilot study aimed to determine if through the use of a protocol with direct questions about whether injuries were caused by battering, the percentage of battered women identified could be increased.

Methods

A retrospective record review of every fourth case of female trauma (excluding motor vehicles and natural disasters) presenting in the MCP ED during calendar year 1976 was conducted (n = 359). The ED studied was located in the inner city and primarily served a lower socioeconomic, Black population. During the first two weeks of 1977, the triage nurses in the ED were trained in administering a protocol aimed at identifying battered women. The protocol contained questions which elicited a trauma history and directly asked if the women were injured by someone. The protocol took approximately 15 minutes to administer. The majority of women responded readily to posed questions in the protocol and appear relieved that someone had directly asked them how they had been hurt. Those who had been battered were given information regarding community resources available to them. The majority of women who had not been battered did not appear to mind being asked if they had been hit or injured by someone. Following a full year of protocol

utilization, the ED records for every fourth case of female trauma during 1977 were reviewed (n = 412). Cases were classified as positive for battering whenever a woman directly stated that her injuries were caused by her having been beaten by someone with whom she was or had been intimately involved. Equivocal cases, e.g., probable and suggestive, were classified as trauma, etiology unspecified.

Results

The percentage of women identified positive for battering increased from 5.6 per cent to 30 per cent following staff training and institution of the protocol in the ED. The difference in proportion was 24.5 per cent (95 per cent confidence intervals: 19.5, 29.5).

With increased detection, an unexpected but impressive finding regarding the relationship between age and battering was noted with 42 per cent of female trauma cases in the 18–20 year age span being battered; 35 per cent of the women 21–30 years of age were also battered as were 18 per cent of injured women 61 years of age and older (Table 1).

Discussion

In an extensive review of the literature, we were unable to find reports of such high percentages of battered women among female trauma cases.[12,16,18–20] While our data do not reflect community incidence of battering nor the incidence of battered women among female trauma patients presenting in emergency departments other than in the one studied, the data do suggest that by directing staff attention to the identification of battered women through the use of a protocol, one can increase the positive identification of injuries caused by battering. While our findings may have reflected an increase in the community prevalence of spousal assault between 1976 and 1977 or a change in the characteristics of the population using the ED, pooled hospital demographic data suggest that the latter did not occur. The question of an increase in the number of battered women presenting to the ED secondary to a shift in prevalence is not so easily dismissed; however, during 1976 and 1977, there were no major changes in the political, cultural, or socioeconomic climate that were of sufficient magnitude to suggest systematic change in prevalence.

TABLE 1—Battering as a Cause of Female Trauma Presenting in an Emergency Department

Age (years)	Total Female Trauma	Positive Cases	% Battered
18–20	47	20	42
21–30	174	61	35
31–40	77	20	26
41–50	45	11	24
51–60	42	7	17
61+	27	5	18
Total	412	124	30

The finding that 42 per cent of female trauma victims between the ages of 18–20 were positively battered is impressive and raises questions regarding the prevalence of battering among teenage girls, a population for whom community services are not available. Finally, the geriatric population is one characterized by frequent injuries subsequent to accidents. With 18 per cent of injured women over 61 years of age admitting that they were injured secondary to being beaten, it becomes essential to increase professional awareness of the risk in this age group.

Unless ED staff maintain a high index of suspicion for battering among female trauma patients, large numbers will go undetected, hence, untreated. Emergency physicians should institute the use of standardized protocols in order to provide adequate care to this population of women.

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Tea and Premenstrual Syndrome in the People's Republic of China

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Abstract: We evaluated the hypothesis that tea consumption causes premenstrual syndrome by studying 188 nursing students and tea factory workers in the People's Republic of China via questionnaires distributed in classes (nursing students) or during routine physical examinations (tea factory workers). Analysis of the data revealed that tea consumption is strongly related to the prevalence of premenstrual syndrome and that the effects are dose-dependent. (*Am J Public Health* 1989; 79:67–69.)

Introduction

The causes of premenstrual syndrome are unknown.[1–3] Putative causes include a hormonal imbalance—particularly a deficit of progesterone,[4] dietary deficiencies,[5–8] a diet rich in highly refined sugar products and/or high fat foods, alcohol use, tobacco use, lack of exercise, oral contraceptive use,[9] and consumption of caffeine-containing beverages.[10]

This study was undertaken to assess the relation between consumption of one type of caffeine-containing beverage (tea) and the prevalence of premenstrual syndrome. To prevent confounding by other (putative) causes of premenstrual syndrome, the study was conducted in a population in which the prevalence of these other causes, including consumption of other caffeine-containing beverages, was expected to be low or zero.

Methods

Study Populations

The study was conducted at three sites in the People's Republic of China: 1) a medical university in Shanghai that trains public health nurses; 2) a nursing school in rural, southern China; and 3) a tea factory in rural, southern China. We expected that the consumption of tea would be less frequent among the nursing students than among the tea factory workers.

The study subjects consisted of all female third and fourth year public health nursing students at the medical university in Shanghai, all female students in one class at the nursing school in rural southern China, and all female workers in the tea factory. The average age of the nursing students was 22.3 years (range, 15 to 33 years), and 24.8 years (range, 17 to 35 years) for the tea factory workers. Participation was 100 per cent for each population.

Data Collection and Analysis

A questionnaire was handed out in class to nursing students at each of the two schools. The questionnaire, although translated into Chinese, was identical in pertinent ways to the questionnaire developed by Rossignol,[10] and solicited information about the presence and severity of premenstrual and menstrual symptoms, demographic characteristics, health-related characteristics (for example, alcohol and tobacco use), use of over-the-counter drugs containing caffeine, and consumption of caffeine-containing beverages. The questionnaires were completed in class and returned.

Data from the tea factory workers were collected during a routine physical examination of each worker. The examinations were performed by female physicians who, during the course of the examination, asked the worker the questions about the presence of menstrual and premenstrual symptoms and about health-related habits including caffeine use. Upon completion of the examination, the physician completed a questionnaire for each worker. The questionnaire was identical to those completed by the nursing students except that the questions related to specific symptoms of premenstrual syndrome were omitted.

We estimated the premenstrual syndrome prevalence odds ratio for two levels of average daily consumption of tea: one-half to four drinks, and four-and-one-half to eight drinks. Area of residence (nursing students, tea factory workers) was controlled by stratifying the data into the two categories of residence and obtaining an estimate for a uniform prevalence odds ratio by the method of maximum likelihood. Heterogeneity of the prevalence odds ratio according to residence was evaluated by the likelihood ratio test proposed by Miettinen, as described by Rothman and Boice.[11] Confidence limits for the prevalence odds ratio were based on Miettinen's test-based procedure,[12,13] using the programs of Rothman and Boice.[11]

The level of severity for premenstrual syndrome was determined by each participant in response to a question asking her to assess the overall severity of her symptoms.

Results

Thirty-nine per cent of the nursing students and 77 per cent of the factory workers reported experiencing premenstrual syndrome. Overall, 74 per cent of these women rated their symptoms as mild, 24 per cent as moderate, and 3 per cent as severe. These proportions were similar in each study group. Among the nursing students, the most common symptoms were breast swelling and tenderness, tiredness, and anxiety, affecting 51, 49, and 34 per cent of women with symptoms, respectively.

Tea was the only reported source of caffeine in both study populations.

Table 1 shows the prevalence of premenstrual syndrome and of putative causes and preventives of premenstrual syndrome by study group. The prevalence of each of the putative causes/preventives, except for tea consumption, was either zero or over 90 per cent for each population. For this reason, any observed association between tea consumption and premenstrual syndrome can not be attributable to confounding by any of these factors. Table 2 shows the prevalence odds ratios of premenstrual syndrome according to average daily tea consumption by residence, and controlling for residence in a high or low tea-consuming region (study group). The prevalence odds ratio for 4.5 to 8 drinks

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TABLE 1—Putative Causes/Preventives of Premenstrual Syndrome for the Study Subjects According to Residence in a High or Low Tea-consuming Region

	Population		
	Low Tea-consuming Region—Nursing Students (124 subjects)	High Tea-consuming Region—Tea Factory Workers (64 Subjects)	Both Populations (188 Subjects)
	%	%	%
Prevalence of Premenstrual Syndrome	39	77	52
Cause/Preventive			
Any Tea Consumption	10	92	38
Caffeine Intake from Cola or Coffee or Over-the-counter drugs	0	0	0
Alcohol Use	0	0	0
Cigarette Use	0	0	0
Daily Exercise (Bike Riding)	over 90	over 90	over 90
Oral Contraceptive Use	0	0	0
Vitamin Supplement Use	0	0	0
Diet	Vegetables, Rice, Pork, No Beef, Few Refined Sugar Products	Vegetables, Rice, Pork, No Beef, Few Refined Sugar Products	Vegetables, Rice, Pork, No Beef, Few Refined Sugar Products

TABLE 2—Prevalence Odds Ratios of Premenstrual Syndrome According to Average Daily Tea Consumption by Residence

Residence (= Study Group)	Average Daily Tea Consumption (Number of Drinks)	Prevalence Odds Ratio of Premenstrual Syndrome (90% Confidence Interval)
Low Tea Consuming (= Nursing Students)	0	1.0
	0.5 to 4	1.4 (0.5, 4.0)
	4.5 to 8	∞ (–)
	0.5 to 8	1.7 (0.6, 4.5)
High Tea Consuming (= Tea Factory Workers)	0	1.0
	0.5 to 4	4.5 (0.9, 21.7)
	4.5 to 8	8.6 (1.7, 42.9)
	0.5 to 8	5.9 (1.4, 25.3)
Both Residences	0	1.0**
	0.5 to 4	2.0 (0.8, 4.7)
	4.5 to 8	9.7 (2.1, 44.5)
	0.5 to 8	2.4 (1.0, 5.6)

**Prevalence odds ratios were estimated controlling for residence in a high or low tea-consuming region (= study group). The $P_{(2)}$-values associated with the likelihood ratio tests evaluating heterogenity of the prevalence odds ratios by residence were 0.29, 0.58, and 0.19 for 0.5 to 4, 4.5 to 8 and 0.5 to 8 drinks of tea per day, respectively. The crude prevalence odds ratios were 3.4 for 0.5 to 4 drinks of tea per day, and 10.0 for 4.5 to 8 drinks of tea per day.

of tea per day were considerably higher than the prevalence odds ratio for 0.5 to 4 drinks per day, consistent with a dose-response relation between tea consumption and premenstrual syndrome.

Discussion

We found a strong association between tea consumption and the prevalence of premenstrual syndrome, particularly for women who consume 4.5 to 8 drinks of tea per day. The effects were similar to the effects reported in a previous study[10] that evaluated the relation between consumption of all caffeine-containing beverages (cola plus coffee plus tea) and the prevalence of premenstrual syndrome. In this previous study, 49 per cent of the caffeine-containing beverages were drinks of cola, 32 per cent were coffee, and only 19 per cent were tea.

In the present study, consumption of tea, especially for the low consumer, typically occurred non-uniformly throughout the week, with pots of tea shared with friends once or twice per week rather than daily. Tea exposure, therefore, varied considerably from day to day, with no exposure on many days and substantially higher exposure (3 to 4 cups of tea) on other days even among low consumers. It is unknown how this pattern of tea consumption may have affected the study findings.

Several factors might account for the observed relation between tea consumption and premenstrual syndrome, including a causal role for total daily fluid consumption (which might be related to average daily tea consumption) or differences in the reproductive histories of the study participants according to the level of tea consumption. None of the commonly suspected, modifiable causes/preventives of pre-

menstrual syndrome, except tea consumption (Table 1), however, is a possible explanation for the study results because none of these other factors varied in frequency according to tea consumption or the presence of premenstrual syndrome. A possible mechanism involving cyclic nucleotides for a causal relation between tea and premenstrual syndrome is provided by Minton.[14,15]

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Differences in the Timeliness of Diagnosis, Breast and Cervical Cancer, San Francisco 1974–85

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Abstract: Cancer registry data for San Francisco (1974–85) were used to identify women at greater risk of late diagnoses for breast and cervical cancers by age and ethnicity. For breast cancer, Black women were at greater risk for late diagnoses. For cervical cancer, women of all ethnic groups ages 50–69 years and Japanese and Filipino women were at greater risk for late diagnoses. (*Am J Public Health* 1989; 79:69–70.)

Introduction

Survival rates for women with early diagnosis of breast and cervical cancer are considerably better than for those with late diagnosis.[1] There is also substantial evidence that mortality from breast cancer for women over age 50 years can be reduced with a combination of regular breast examination and mammography.[2–4] For women ages 20–70 years there is good evidence, although no randomized controlled trials have been done, that cervical cancer mortality can be reduced by regular cervical pap smears.[5–7]

Despite this, the utilization of screening tests is low. A 1983 national survey found that only 45 per cent and 14 per cent of women ages 50–70 years had had annual physical examinations and mammograms, respectively. Among women ages 20–39, 79 per cent had had pap smears every three years, compared to only 57 per cent of women ages 40–70 years.[8]

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Methods

The study uses data collected on San Francisco women diagnosed with breast cancer (ICD-0 1976 Code 174) or cervical cancer (ICD-0 1976 Code 180) between 1974 and 1985 by the San Francisco-Oakland SMSA cancer registry, which identifies approximately 98 per cent of incident cases of cancer occurring in five San Francisco Bay Area counties.[9] The purpose of this analysis was to examine differences in the timeliness of diagnosis of female breast and cervical cancer in different age and ethnic groups.

Based on the ability of screening programs to detect high percentages of breast cancers without nodal involvement,[4] breast cancers were classified as early (in situ and localized stages) or late (regional or remote stages). As regular pap smears can detect a high percentage of cervical cancers before these became invasive,[5] cervical cancers were classified as early (in situ) or late (localized, regional or remote).

Distribution of the timeliness of diagnosis (early and late) by age group (<30, 30–39, 40–49, 50–59, 60–69 years) for these cancer types were calculated for White non-Hispanic, White Hispanic (primarily Mexican American), Black, Chinese, Japanese and Filipino women for the years 1974–85. The upper age limit of 70 years was chosen to follow the National Cancer Institute recommendations of screening for these cancers under the age of 70 years.[10]

Chi square tests for trend were used to examine the association between late diagnosis of breast and cervical cancer with age. To examine ethnic differences in late diagnoses for these cancers odds ratios (with 95% confidence intervals) were calculated for each of the other ethnic groups using White non-Hispanic women as the reference group.

Results

Breast Cancer

We found that 1,568 (43.5 per cent) of the 3,608 breast

cancers diagnosed in women less than 70 years of age were late diagnoses (regional or remote tumors) with no differences in the percentage of late diagnoses by age (range 40.9–44.4 per cent Chi square for trend = 0.15, p = 0.70); this was true for all ethnic groups. Most cases of breast cancer (69.5 per cent) were diagnosed in women ages 50–69 years and therefore most (69.1 per cent) late diagnoses were in this age group. Of the 1,568 late diagnoses, 1,094 (70 per cent) occurred in White non-Hispanic women. Among the ethnic groups, Black women had significantly more late diagnoses than White non-Hispanic women (Table 1). One hundred twenty-eight (64.3 per cent) of the 199 late diagnoses in Black women occurred in women ages 50–69 years.

Cervical Cancer

Of the 2,090 cervical cancers diagnosed in women less than 70 years old, 430 (20.6 per cent) were late diagnoses (localized, regional or remote tumors). In contrast to breast cancer, there is a marked increase with age in the percentage of women with late diagnoses (Table 2). The odds ratio for 60–69 year-old women compared with women less than 30 years old was 20.2 (95 per cent CI = 12.8, 32.0). Two hundred nine (48.6 per cent) of the 430 late diagnoses occurred in White non-Hispanic women. To examine the association between late diagnosis and ethnicity for cervical cancer, directly weighted pooled estimates of odds ratio (and 95 per cent CI) were calculated to adjust for the confounding effect of age.[11] Of the ethnic groups, Japanese and Filipino women had significantly more late diagnoses than White non-Hispanic women (Table 3).

Discussion

Late diagnosis for breast cancer was common for all age groups. However, 69.1 per cent of late diagnoses were in women ≥50 years of age reflecting the higher incidence of breast cancer in these age groups.[12] Black women with breast cancer tend to be from poorer social classes than White women with breast cancer.[13] The later diagnosis of breast cancer in Black women may reflect lack of access to screening mammography services. The National Cancer Institute has recommended that by the year 2000 80 per cent of women aged 50–59 years[10] should have annual breast examination and mammography. Black women 50–69 years of age should be a high priority for screening.

The higher percentage of late diagnoses for cervical cancer among older women may be due to lower cervical pap smear rates in older women.[8] This in turn may be due to the erroneous perception that cervical cancer is largely a disease of younger women. Cervical cancer incidence rates for San Francisco-Oakland SMSA for women 60–69 years of age are, in fact, higher than for all younger age groups.[12] The National Cancer Institute recommends that by the year 2000 90 per cent of women 20–39 years of age and 80 per cent of women 40–69 years of age should have a cervical pap smear every three years. Efforts in this regard in San Francisco should be directed particularly toward women 50 years of age and older and also toward Japanese and Filipino women.

TABLE 1—Odds Ratios by Ethnic Group, Female Breast Cancer, San Francisco, 1974–85

Ethnic Group	Odds Ratio*	95% CI
White Hispanic	0.98	0.74, 1.30
Black	1.71	1.36, 2.14
Chinese	0.86	0.66, 1.12
Japanese	1.16	0.72, 1.84
Filipino	0.97	0.66, 1.41

*Reference group: White non-Hispanic

TABLE 2—Stage of Diagnoses by Age Group, Cervical Cancer, San Francisco, 1974–85

Age (years)	Stage		Total	
	Early	Late	No.	(%)
<30	599 (94.9)*	32 (5.1)	631	(30.2)
30–39	633 (87.1)	94 (12.9)	727	(34.8)
40–49	233 (72.8)	87 (27.2)	320	(15.3)
50–59	107 (46.7)	122 (53.3)	229	(11.0)
60–69	88 (48.1)	95 (51.9)	183	(8.8)
Total	1660 (79.4)	430 (20.6)	2090	(100.0)

*Row percentage
Chi square for trend = 362.1 p = < 0.000001

TABLE 3—Odds Ratios by Ethnic Group, Cervical Cancer, San Francisco, 1974–85

Ethnic Group	Odds Ratio*	95% CI
White Hispanic	1.31	0.92, 1.87
Black	1.07	0.76, 1.49
Chinese	0.97	0.64, 1.46
Japanese	2.65	1.21, 5.83
Filipino	2.17	1.30, 3.62

*Reference group White non-Hispanic

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Breast Cancer Incidence in Young Women by Estrogen Receptor Status and Race

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Abstract: A population-based study was utilized to calculate breast cancer incidence rates in White and Black women, ages 30 to 54, according to tumor estrogen receptor status. Both racial groups had higher incidence curves for estrogen receptor negative breast cancer between ages 30 and 49. There was an excess of receptor negative cancer in young Black women, an observation that may help explain the racial disparity in breast cancer survival. (*Am J Public Health* 1989; 79:71–73.)

Introduction

Clinical studies suggest that estrogen receptor analysis should be performed on all primary breast cancers because results provide valuable therapeutic and prognostic information.[1] The presence (estrogen receptor positive) or absence (estrogen receptor negative) of specific estrogen binding protein in breast cancer is related to the biologic characteristics of the tumor. Receptor negativity is associated with larger tumor size and more rapidly proliferating tumor tissue.[2,3] Estrogen receptor status also varies according to some breast cancer risk factors, notably age, menstrual status, and race.[4–11] Despite evidence that receptor results are correlated with prognostic and risk variables, we are aware of only one study that has examined the incidence of breast cancer stratified by receptor results and that study was limited to a predominantly Caucasian population.[12] The present report provides incidence rates of estrogen receptor positive and estrogen receptor negative breast cancer separately for White and Black women based on data collected by a population-based cancer registry.

Methods

Data from the Georgia Center for Cancer Statistics, an affiliate of the Surveillance, Epidemiology and End Results (SEER) Program of the National Cancer Institute, were used to estimate breast cancer incidence rates according to estrogen receptor status. All cases of newly diagnosed, pathologically confirmed breast cancer among women 30–54 years of age were identified during the period of December 1980 through December 1982. Estrogen receptor assay results, abstracted from medical records, were available on 82.4 per cent of the 637 breast cancer cases ascertained during the two-year period (Table 1). Overall, two-thirds of the tumor specimens were assayed at the same laboratory, and 98.6 per cent of the samples were assayed by the standard multipoint titration method using dextran-coated charcoal.[7]

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TABLE 1—Number of Incident Breast Cancer Cases Diagnosed by Age and Estrogen Receptor Status, Atlanta, Georgia, 1980–82

Age (years)	Total No. Cases	Estrogen Receptor Status			
		Positive	Negative	Borderline	Unknown
30–34	55	15	33	2	5
35–39	114	36	57	1	20
40–44	133	45	62	5	21
45–49	174	58	69	6	41
50–54	161	77	54	5	25
Total	637	231	275	19	112

There were 231 estrogen receptor positive, 275 estrogen receptor negative, and 19 estrogen receptor borderline (i.e., equivocal results) breast cancers. (Estrogen receptor borderline tumors behave clinically like receptor negative tumors[1,2]; these 19 cases were therefore combined with the estrogen receptor negative group.) Women of other races (n = 3) and those with unknown receptor values (n = 112) were excluded from the calculation of incidence rates. The percentage of women with unknown receptor values did not differ by race: Whites, 17.9 per cent; Blacks, 16.9 per cent. Within each race, the percentage of unknown receptor values was greatest for women ages 45–49 years: Whites, 22.5 percent; Blacks, 28.6 per cent; the racial distribution of cases with unknown receptor results (Whites, 78.6 per cent; Blacks, 21.4 per cent) was similar to that of cases included in the analysis (Whites, 77.6 per cent; Blacks, 22.4 per cent). Population estimates for age and race were obtained from the US Census, and are specific for the five metropolitan counties served by the Atlanta cancer registry.[13]

Results

The overall breast cancer incidence rates increased with age. Rates of estrogen receptor negativity were higher than rates of estrogen receptor positivity in each five-year age group up to the ages of 50–54 years, where the curves crossed, a pattern generally consistent for Whites (Figure 1) and Blacks (Figure 2). However, the differential between curves for Black women was substantially wider than that for Whites, and Black women ages 30–44 years had higher incidence rates of estrogen receptor negative breast cancer than Whites.

Incidence rate ratios were calculated by age and receptor status (Table 2). Although most of the 95 per cent confidence intervals (CI) include unity, Whites generally have higher receptor positive cancer rates than Blacks among the age groups 35–54 years; Black women ages 30–44 had an excess of estrogen receptor negative cancer.

Logistic regression[14] was used to evaluate the potential confounding effects of stage of disease and laboratory variability on differences in receptor status between Whites and Blacks. The overall age-adjusted risk estimate for an estrogen receptor negative (as opposed to an estrogen receptor positive) tumor in Blacks compared to Whites was 1.9 (95 per cent

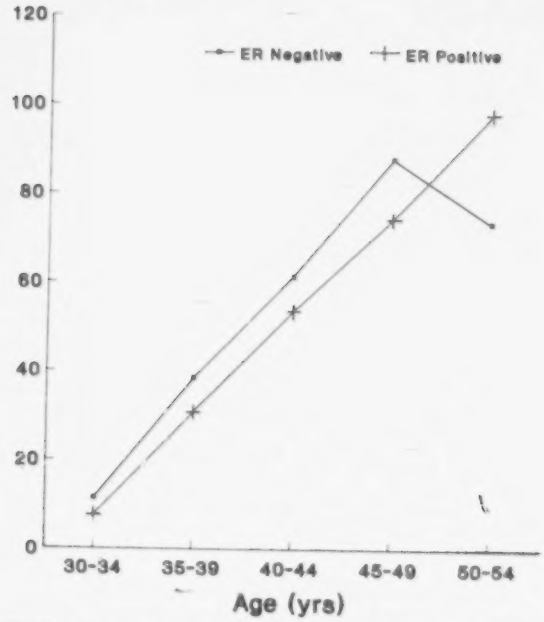

FIGURE 1—Average Annual Age-Specific Breast Cancer Incidence per 100,000 White Women

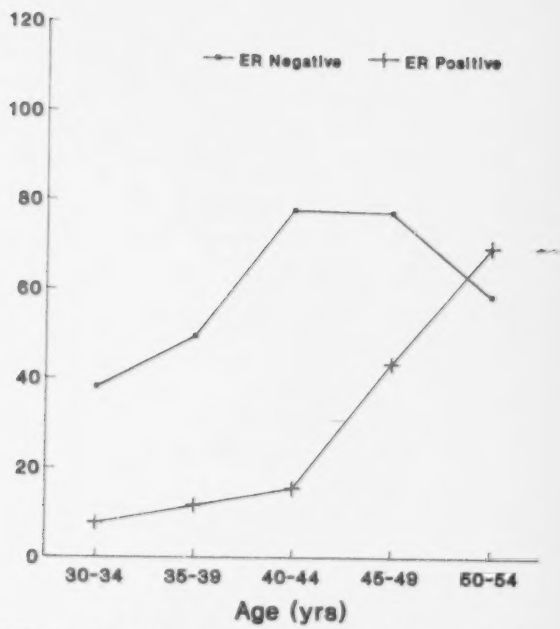

FIGURE 2—Average Annual Age-Specific Breast Cancer Incidence per 100,000 Black Women

CI = 1.2, 3.1) and did not change materially after adjustment for stage of disease at diagnosis or laboratory where the receptor assay was performed.

Discussion

The results of this study may have been influenced by a number of factors. Estrogen receptor results were unavailable for 17.6 per cent of the cases identified for study, but the age and race distributions of these cases did not differ substantially from those of cases included in the analysis. Exclusion of women with unknown receptor values resulted in slightly lower age-specific incidence rates for Whites and Blacks. If women with unknown receptor status had a distribution of estrogen receptor results similar to those with known status, the age-specific incidence patterns would be similar to those presented here.

Laboratory variability was not a factor. The majority of tumor specimens (for both Whites and Blacks) in this series of cases were analyzed at the same laboratory utilizing the same methodology. Furthermore, based on multivariate analyses, Blacks had a higher risk estimate than Whites for estrogen receptor negative (versus receptor positive) cancer, controlling for laboratory.

Multivariate analyses showed a higher risk estimate for estrogen receptor negative cancer (as opposed to estrogen receptor positive cancer) in Blacks compared to Whites after adjusting for age and clinical stage at diagnosis. Thus it does not appear that differences in extent of disease explain the observed differences in White:Black breast cancer rates by age and receptor status.

These age- and race-specific patterns of incidence rates for breast cancer are consistent with prior clinical and epidemiologic observations of an increase in the proportion of estrogen receptor positive cancer with advancing age,[4,5,10,12] and a lower prevalence of estrogen receptor positive breast cancer among Black women compared to White women.[5,8–10] The data suggest that younger women and Black women may be more likely to develop a more aggressive form of breast cancer, i.e., estrogen receptor negative disease.

Estrogen receptor negativity is associated with shorter disease-free intervals and survival in some cases.[1,2,15] In addition, Black women apparently have a lower proportion of receptor positive breast cancer compared to their White counterparts, in pre- and postmenopausal groups,[6,10,16] partly because they have more advanced disease at diagnosis and a higher proportion of poorly differentiated tumors.[6,16,17] The present analysis suggests that Blacks are at higher risk of estrogen receptor negative (as opposed to receptor positive) breast cancer than Whites, after adjustment for age and stage of disease at diagnosis.

Racial variation in estrogen receptor results has been

TABLE 2—White:Black Incidence Rate Ratios* by Age and Estrogen Receptor Status, Atlanta, Georgia, 1980–82

	Estrogen Receptor Status			
	Positive		Negative	
Age (years)	Point Estimate	(95% CI)**	Point Estimate	(95% CI)**
30–34	0.99	(0.31, 3.17)	0.30	(0.16, 0.56)
35–39	2.63	(0.97, 7.16)	0.77	(0.44, 1.37)
40–44	3.43	(1.31, 9.00)	0.79	(0.47, 1.33)
45–49	1.71	(0.85, 3.44)	1.14	(0.65, 1.98)
50–54	1.41	(0.78, 2.55)	1.25	(0.65, 2.40)

*Rate per 100,000 women per year in Whites divided by rate in Blacks.
**Test-based 95% confidence intervals.

hypothesized as one explanation for the observed Black-White differences in breast cancer survival.[6,8,16,18] The present data provide further evidence that receptor negative breast cancer in Black women may be a contributing factor to their survival disadvantage. Future studies of Black-White differences in breast cancer survival should account for tumor estrogen receptor status.

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Bowel Function and Breast Cancer in US Women

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Abstract: We studied bowel function in relation to 123 breast cancer cases among 7,702 women from the US NHANES I Epidemiologic Follow-up Study. Results suggest a slight increased risk of breast cancer for both decreased frequency of bowel movements (relative risk = 1.5, 95% confidence interval = 0.8, 2.7) and firm stool consistency (RR = 1.8, 95% CI = 1.0, 3.2.) These observations are consistent with an hypothesized association between constipation and increased risk of breast cancer. (*Am J Public Health* 1989; 79: 73–75.)

Introduction

There is indirect evidence that bowel function may be related to breast cancer risk. Petrakis and King found severe constipation (fewer than three bowel movements per week) to be associated with cytologic abnormalities in epithelial cells from breast fluid,[1] which are related to epithelial dysplasia in breast tissue.[2] Atypical proliferative breast disease is a significant risk factor for breast cancer in women; we, and others, have shown the risk to increase with increasing degree of epithelial atypia.[3,4]

Putative mutagens and carcinogens have also been detected in nipple aspirates of breast fluid.[5] Although the precise origin and mechanism of carcinogen delivery to the breast tissue are unknown, involvement of fecal mutagens, bowel function, and the enterohepatic circulation has been hypothesized. Intestinal bacteria produce carcinogens and mutagens, presumably through their actions on dietary constituents and/or bile acids.[6–8] Breast secretory (apocrine) epithelia selectively absorb and concentrate substances from the circulation originating from the gastrointestinal tract.[1,6,9]

Constipation, which results in greater contact time of stool in the intestine and hard stool consistency, may increase formation and absorption of fecal mutagens into the enterohepatic circulation and delivery to breast tissue.[10,11] We describe the first study relating aspects of bowel function directly to the risk of breast cancer in a cohort of women in the first National Health and Nutrition Examination Survey (NHANES I) Epidemiologic Follow-up Study.

Methods

NHANES I and its augmentation survey were conducted by the National Center for Health Statistics (NCHS) from 1971 to 1975.[12,13] These surveys provided cross-sectional information on medical history, anthropometric, biochemical, clinical, demographic, and nutritional factors in a large sample selected to represent the non-institutionalized popu-

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TABLE 1—Adjusted Relative Risks (RR) for Breast Cancer in Relation to Bowel Function

Bowel Function	Cases (N)	Non-cases (N)	Age-adj. RR	95% CI	Multi-adj.[a] RR	95% CI
Bowel problems						
Diarrhea	3	330	0.5	0.2, 1.7	0.6	0.2, 1.9
No Trouble	89	5642	(1.0)	—	—	—
Constipation	31	1597	1.0	0.7, 1.5	1.1	0.7, 1.6
Stool Frequency						
>1×Day	12	698	1.1	0.6, 2.0	1.1	0.6, 2.0
1×Day	79	5257	(1.0)	—	—	—
4–6×Week	15	801	1.3	0.7, 2.2	1.3	0.7, 2.3
≤4×Week	16	841	1.4	0.8, 2.4	1.5	0.8, 2.7
Stool Consistency						
Loose	4	266	1.3	0.5, 3.5	1.3	0.5, 3.6
Normal	53	4977	(1.0)	—	—	—
Firm	18	840	1.7	1.0, 2.9	1.8	1.0, 3.2

[a]adjusted for current age, age at menarche, age at first birth, age at menopause, body mass index, socioeconomic status and family history

lation of the US. Children ages 1 to 5 years, women ages 20–44 years, the elderly (aged 65 years and over), and low-income individuals were oversampled. Details of a follow-up study of 14,407 adults (25 years or older) conducted between May 1982 and August 1984 were provided by Huntley, *et al.*[14] Of the original participants, 12,554 were successfully traced and reinterviewed (14 percent by proxy). Follow-up data included additional interview information, weight and blood pressure measurements, hospital and nursing home records, and death certificate information.

One hundred twenty-three (123) verified cases of breast cancer were identified in the NHANES I follow-up cohort. Women with self-reported breast cancer only (N = 36) were not considered as cases and conservatively included in the control group. There were 7,643 women in the cohort who did not develop breast cancer.

During the baseline interview, participants had been asked several questions about usual bowel habits: usual frequency (once per week or less, two to three times per week, four to six times per week, once per day, two to three times per day, or four or more times per day); characterization of frequent "bowel problems" (no trouble, frequent constipation or diarrhea, or both), and frequency of laxative use. Information regarding usual consistency of the stool (normal, loose, firm, or mixed) was asked only at follow-up.

Adjusted relative risks of breast cancer were calculated for bowel problems, bowel frequency, stool consistency, and laxative use using proportional hazards models from the SAS statistical procedure[15] PHLGM. Risk factors for breast cancer included in the multivariate models were current age as a continuous variable, age at menarche (<12, 12–13, ≥14), age at birth of the first child (<21, ≥22, nulliparous),* age at menopause, (≤45, >45, premenopausal), body mass index (<30.0, ≥30.0), socioeconomic status as poverty index ratio (<3.75, ≥3.75), and family history (one or more first-degree relatives with breast cancer; negative or positive). Reference categories for each of these variables are the first ones listed in each case. Missing data indicators were included in the Cox regression models in order to retain cases with data missing for specific variables only. Weight/stature[1.5] was employed as the body mass index in this study due to its high correlation with weight and independence from stature across age groups in this population sample.[17] Body mass index was chosen as a reliable index of body size that is of proven interest as a risk factor for breast cancer, and has strong significance for the complex inter-relations between nutrition, body size, and breast cancer.[10,18]

*Young age at first birth is highly correlated with increased parity, and the association of high parity with low risk of breast cancer has been explained by its association with young age at birth of the first child.[16]

Results

Laxative use was not a significant risk factor for breast cancer in the cohort (RR = 1.03, 95% CF = 0.91, 1.12) and it was not included in the final models. Table 1 reports relative risks of breast cancer for bowel problems, stool frequency, and stool consistency, adjusted for age and multiple risk factors. While a history of frequent diarrhea appeared to be associated with lower risk (RR = 0.5), the confidence interval was wide, based upon this number of cases. Self-reported constipation was not a predictor of adjusted breast cancer risk. However, the relative risk of breast cancer appeared to increase with decreased frequency of bowel movements. Firm consistency of stool was also associated with increased risk of breast cancer (RR = 1.8, 95% CI = 1.0, 3.2). Including multiple breast cancer risk factors (in addition to age) into the models somewhat increased the strength of the associations between breast cancer risk and both decreased stool frequency, and firm stool consistency. Excluding the 36 cases of non-verified breast cancer from the analysis, rather than including them in the control group, did not alter these results.

Discussion

Taken together these results suggest a slightly increased risk of breast cancer in women with less frequent and firm stools. Although these differences were not all "statistically significant" at $p < 0.05$, relative risks were at or above 1.5 with lower limits of the confidence intervals close to or including 1.0. These observations are consistent with a model of bowel function, breast and apocrine gland function, and breast cancer described by Petrakis, *et al*,[1,9] and recently critically reviewed.[10,11] However, our analyses were based upon self-reported bowel habits which could not be defined in totally objective terms. The self-reporting of "bowel problems" was not quantified in terms of stool frequency and consistency and is probably the least reliable measure,

perhaps accounting for the absence of risk associated with self-reported constipation.

Self-reporting of other bowel habits may also contribute to subjective misclassification of participants which would operate against the observation of any true relations between bowel function and breast cancer that may exist. Other limitations of these data include the fact that self-reporting of stool consistency was obtained at follow-up, rather than at baseline, and that other breast cancer risk factors not included in the model could have been confounders.

Further investigations into the relations between breast and apocrine gland function, bowel function and pre-neoplastic and neoplastic breast conditions are necessary to confirm the observed associations, and to elucidate the etiologic components of this model of breast carcinogenesis. Such investigations should focus on collection of dietary information and biologic samples (blood, breast fluid, stool), together with information on established breast cancer risk factors, in a defined population of women.

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X-Ray Examinations during Pregnancy: National Natality Surveys, 1963 and 1980

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Abstract: Based on 1963 and 1980 National Natality Surveys, the rate of medical x-ray examinations during pregnancy per 100 mothers fell 34.2 percent. A decrease in chest x-ray examinations accounted for almost all of the decline in total x-ray examinations. The reductions were greater for older mothers and those who were not White. While the number of births fell from 4,071,000 in 1963 to 3,612,000 in 1980, the number of pelvimetry examinations actually increased by 45,000. (*Am J Public Health* 1989; 79:75–77.)

Introduction

The sensitivity of the fetus to ionizing radiation has been well documented.[1–4] The per rad increased risk of childhood leukemia has been estimated to be between 1.3 and 2.8 times the national incidence.[1–2] This increase is substantially larger than the increase expected from an equivalent dose administered to an adult population.[3]

Given the increased awareness during the 1960s and 1970s of the danger of fetal exposure to ionizing radiation, many observers predicted a decline in fetal dose from medical diagnostic procedures that utilize ionizing radiation.[5] Ultrasound became increasingly available during the 1970s, and was widely believed to have reduced fetal dose.[6] The precise health effects of fetal exposure to ultrasound are unknown, but some animal studies have suggested that the fetus may be adversely affected.[7]

We utilized data from the 1963 and 1980 National Natality Surveys (NNS) to obtain information about exposure to ionizing radiation among mothers who had live births.[8]

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Methods

The National Center for Health Statistics sampled 9,941 certificates of live birth in the 1980 NNS and 4,096 certificates of live birth in the 1963 NNS. Radiation information was sought from hospitals where mothers had delivered, from attendants at delivery, and from medical providers of radiation for all births. In both the 1963 and 1980 NNSs, detailed information regarding the mother's exposure to ionizing radiation during the 12 months prior to delivery was sought from the hospitals and from the attendants at delivery. The hospitals, attendants at delivery, and married mothers were all asked to identify any other sources that might have provided the mother with x-ray examinations prior to delivery. Health care providers identified in this way were also sent a questionnaire concerning the mother's radiation exposure.

Post-stratified ratio estimation procedures were used to produce national estimates of mothers exposed to radiation. These estimates of exposure may be understated to the extent that respondents failed to identify all health care providers that the mother saw during pregnancy, and to the extent that sources contacted failed to report all procedures performed. Although information was collected about radiation exposure during the 12 months before delivery, this analysis is limited to procedures performed during pregnancy. For relevant rates, 95 per cent confidence intervals were calculated.

In the 1963 NNS, response rates were as follows: married mothers—86 per cent; hospitals—98 per cent; and attendants at delivery—93 per cent.[9] In the 1980 NNS, response rates were as follows: married mothers—80 per cent; hospitals—76 per cent; and attendants at delivery—62 per cent.[6,7] The response rates for hospitals and attendants at delivery in the 1980 NNS understated the proportion of cases for which radiation exposure information is available since this information was collected not only from hospitals and attendants at delivery, but also from other providers as described above.

Results

The percentage of mothers who had one or more x-ray examinations during their pregnancies decreased in all age groups in 1980 (Table 1). Mothers 30–34 years of age experienced the largest decline (43.5 per cent), followed closely by mothers 35 years and over (41.6 per cent). All regions experienced a decline. The decline was greater for races other than White. Table 2 illustrates the trends in number and rate of x-ray examinations by broad classes of examinations. The total number of all x-ray examinations decreased by approximately 462,000, a 34.2 per cent decline in the rate per 100 mothers. The total number of chest x-ray exams decreased by approximately 452,000. However, the number of pelvimetries increased by 45,000, from 5.1 to 7.0 per 100 mothers.

Discussion

Other studies suggest a decline in x-ray examinations during pregnancy, but do not address whether a resultant decrease in fetal dose had occurred. Our analysis shows that almost all of the decrease in total examinations can be attributed to a decline in chest x-ray examinations. The 452,000 fewer chest x-ray exams performed in 1980 closely mirrors the drop of 462,000 total examinations noted between 1963 and 1980.

Studies during the 1970s questioned the value of routine prenatal chest x-rays and discouraged pelvimetry,[10–12] and so a drop in x-ray examinations during pregnancy was to be expected. For example, only a single benign lesion was uncovered by routine prenatal chest x-rays in a study of 1,200 patients.[10] Another study of 11,000 consecutive routine prenatal chest x-rays found no new useful information.[13]

On the other hand, the rate of pelvimetry per 100 mothers increased from 5.1 in 1963 to 7.0 in 1980, an increase of 37.0 per cent. The pronounced increase in the cesarean section rate that occurred between 1965 and 1980[14] may have contributed to the increase in the pelvimetry rate.

The absorbed fetal dose from a chest x-ray is relatively low.[15] The dose from a frontal projection chest x-ray has been estimated to be only approximately 20 millirad.[16] The lungs are relatively permeable to x-rays because air is less radiodense than soft tissue. Additionally, the fetus is not in the radiographic field during a chest x-ray and receives only scatter radiation. The estimated fetal dose from a posterior-anterior chest x-ray is in the range of 1 millirad.[13] Consequently, the degree of decreased fetal dose from the decline in chest x-ray examinations is relatively small.

In contrast to chest x-rays, the fetus is in the radiographic field during pelvimetry. This fact coupled with the greater radiodensity of the abdomen compared to the thorax results in an estimated fetal average dose of 620 millirad from a single pelvimetry,[5] far greater than the dose from a chest x-ray.

The decline from 1,117,000 to 655,000 in total examinations during pregnancy utilizing ionizing radiation has not translated into a commensurate decline in fetal dose. The large drop in chest x-rays has contributed little to decreasing fetal dose. Pelvimetry, which contributes much more to the fetal dose has actually risen in both absolute and relative terms between 1963 and 1980, according to the National Natality Survey (NNS).

The use of prenatal ultrasound is increasing[17] and has the

TABLE 1—Rate of Medical X-ray Examinations during Pregnancy per 100 Mothers by Selected Characteristics of Mothers: National Natality Surveys, 1963 and 1980

Characteristics of Mothers	1963 Rate (95% CI)	1980 Rate (95% CI)	Per Cent Change in Rate
Total	27.5 (26.5, 28.5)	18.1 (17.3, 18.9)	−34.2
Age (years)			
<20	28.4 (25.7, 31.1)	22.7 (20.4, 25.0)	−20.1
20–24	29.0 (27.2, 30.8)	17.9 (16.4, 19.4)	−38.3
25–29	23.8 (21.9, 25.7)	17.4 (16.2, 18.6)	−26.9
30–34	27.8 (25.1, 30.5)	15.7 (13.8, 17.6)	−43.5
35>	29.6 (26.4, 32.8)	17.3 (13.1, 21.5)	−41.6
Race			
White	26.2 (25.1, 27.3)	18.8 (17.7, 19.9)	−28.2
Other	34.6 (31.9, 37.3)	15.1 (13.6, 16.6)	−56.4
Region of Residence			
Northeast	26.2 (24.1, 28.3)	15.8 (13.7, 17.9)	−39.7
Midwest	28.7 (26.7, 30.7)	21.7 (19.7, 23.7)	−24.4
South	25.6 (23.6, 27.4)	18.8 (17.0, 20.6)	−26.6
West	31.0 (28.4, 33.6)	14.4 (12.6, 16.2)	−53.6

TABLE 2—Number of X-ray Examinations and Rate of Examinations during Pregnancy per 100 Mothers by Type of Examination: National Natality Surveys, 1963 and 1980

		Abdominal x-rays				
	Total x-rays	Pelvimetry	Other not related to pregnancy	Other related to pregnancy	Chest x-rays	Other x-rays
1963						
Number of Examinations	1,117,000	209,000	158,000	47,000	648,000	56,000
Rate per 100 Mothers[a]	27.5	5.1	3.9	1.2	15.9	1.4
95% CI	(26.5, 28.5)	(4.6, 5.6)	(3.5, 4.3)	(1.0, 1.4)	(15.1, 16.7)	(1.1, 1.7)
1980						
Number of Examinations	655,000	254,000	129,000	45,000	196,000	31,000
Rate per 100 Mothers[a]	18.1	7.0	3.6	1.3	5.4	0.9
95% CI	(17.3, 18.9)	(6.5, 7.5)	(3.2, 4.0)	(1.1, 1.5)	(4.9, 5.9)	(0.7, 1.1)
Percent 1963 and 1980 Change in Rate	−34.2	+37.0	−8.0	+7.9	−65.9	−37.6

[a]Based on 4,071,000 mothers in 1963 and 3,612,000 mothers in 1980.

potential to replace many x-ray examinations during pregnancy,[5] therefore a further decline in total x-ray examinations is anticipated. The 1988 National Maternal and Infant Health Survey (NMIHS), which includes information on x-ray examinations to 10,000 mothers of live births will give us our next glimpse as to whether total x-ray examinations will decline further, encompassing a decrease in both chest x-ray examinations and pelvimetry.[18]

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Records, Recall Loss, and Recall Bias in Pregnancy: A Comparison of Interview and Medical Records Data of Pregnant and Postnatal Women

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Abstract: To determine the agreement between interview-based and medical records data concerning illnesses and medications early in pregnancy, we compared the interviews of 202 women with the ongoing records collected during their pregnancies. Substantial underrecording of most transient illnesses was found. Fewer short-term illnesses were reported by postpartum women than still-pregnant women, suggesting the potential for recall bias or loss. (*Am J Public Health* 1989; 79:78–80.)

Introduction

Collection of data about the occurrence and timing of pregnancy events is fraught with potential problems in reliability. If a case and control group are compared, recall bias occurs if there is differential reporting by mothers of affected infants versus those of "healthy babies." Although a standard medical record, collected prospectively throughout pregnancy, may yield unbiased data, a recent comparison of medical records and women's reports of intrapartum events questioned whether the medical records are accurate.[1]

This paper compares self-reported data from still-pregnant and postpartum women with records data on potentially important events which occur during early pregnancy: medication use and history of transient illnesses.

Methods

The data presented derive from a sample of women from two control groups used in a larger study of risk factors for spontaneous abortion.[2] Three hospitals participated in the study between January 1984 and April 1985; these hospitals were the only three hospitals in a city of over 600,000 people which provided both emergency and routine obstetrical care. Overall response rate was 87.4 per cent for contacted prenatal women and 87.5 per cent for postnatal women in the larger study. In this report, the postnatal group consisted of women still in one of the three study hospitals following delivery of a normal infant. The prenatal group consisted of women less than 25 weeks pregnant, whose names were drawn from prenatal case lists provided by physicians admitting to one of the study hospitals, and matched by age (within five-year intervals) and parity (0, 1, 2+). "Late attenders" are thus not represented in this sample; 94.6 per cent of our entire group attended for care before the 16th week of gestation.

All women were interviewed regarding specific periconceptual exposures occurring up to four months following the last menstrual period. In addition to direct open-ended questions regarding exposure to short-term illnesses and medications, probes were used as "examples." For short-term illnesses, the probes included "colds, the flu, bladder infections, measles, mumps, other infections, etc.;" for medications the probes included "prescriptions, aspirin, cough medications, antinausea pills, cold medications, allergy preparations, vitamins, laxatives, etc."

At the end of the study, an abstractor who was unaware of the information on the original interview searched the hospital records of 101 pairs of prenatal and postnatal women. All hospital records contain a copy of the standard outpatient prenatal record form, which is filled out on an ongoing basis throughout the pregnancy by the attending physician and submitted to the hospital before delivery. This form contains information on past medical and drug history, and a record of each prenatal visit and problems discussed therein. Specific questions about vaginal bleeding and nausea appear on the form. For the women still pregnant at time of interview, prenatal forms were accessed after their delivery. However, only entries which covered the time up to and including the interview date were included in the analysis. Concordance was recorded only if interview and chart information agreed in the number and types of reports. Broad classifications of illnesses or drugs were used and exact concordance of proprietary names was not required.

We computed kappa values to compare the data abstracted from the prenatal hospital records with the interview data.[3] We also compared the 71 pregnant women who were interviewed after the sixteenth week of pregnancy, to their matched postnatal controls, calculating the differences in proportions and 95 per cent confidence intervals.[4] A matched analysis using binomial probabilities yielded similar findings and is not presented here.

Results

Health complaints were common during the first four months of pregnancy, but agreement between interview data and chart data was poor, with no kappa values exceeding 0.24 (Table 1). The major portion of agreement in all cases came from the absence of a complaint on either data source; positive concordance was much less common.

The interviews revealed a higher proportion of potentially infectious or febrile events than did the charts, but the charts reported more vaginal bleeding and nausea. For medication use, crude agreement was highest for prescription drugs.

When the interview histories of the first four months of pregnancy, as given by pregnant and postnatal women are compared, a pattern of more complete reporting of nausea/vomiting, upper respiratory infections, and flu/febrile illnesses by still-pregnant women is seen (Table 2). The pooled data of all short-term illnesses indicates more overall reports by still-pregnant women. For medication use, however, the difference between exposure reports by pregnant versus postnatal women is inconsequential in all categories.

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TABLE 1—Agreement between Chart and Interview for Short-term Illness and Medication Reporting

	Source of Report						
	Interview Only	Chart Only	Both	Neither	Crude* Agreement	kappa¶	Interview: Chart Ratio†
Short-term Illness							
Upper respiratory infection/symptoms	45	2	3	152	76.7	.07	9.60
"Flu-like" or febrile illness	25	2	4	171	86.6	.19	4.83
Nausea/vomiting	8	34	5	155	79.2	.11	0.33
Vaginal spotting/ bleeding	2	26	3	171	86.1	.14	0.17
Other complaints	7	13	4	178	90.1	.24	0.65
Medication Type							
Prescription drugs	12	10	13	167	88.6	.48	1.09
Over-the-counter drugs	65	4	3	130	65.8	.02	9.71
Vitamins/supplements	88	11	82	21	51.0	.07	1.83

*Crude Agreement = [number of instances of reports on both records + number of instances of reports on neither record]/total number studied.
¶Kappa value is a measure of agreement beyond that which would be expected to occur by chance. A maximum value for kappa is 1.0; levels below 0.4 are generally taken to indicate poor agreement.
†Interview: Chart Ratio = number of instances of positive reports on interview/number of instances of positive reports on chart.

Discussion

It has been noted that a high proportion of women interviewed one to 15 months postpartum fail to give drug and intermittent illness history identical to that recorded in an interview at five months' gestation.[5] Agreement between birth certificates and mother's reports of time of onset of prenatal care is concordant for month in only 44 per cent of cases,* and concordant for trimester of onset in 88 per cent for White subjects,[7] with mothers tending to report earlier care than the certificates. Distant recall of drug use in past pregnancies has been found to be poor even for prescription drugs.[7] Some have suggested that, because this recall error (i.e., loss) can create inaccurate results in retrospective studies, "studies of past discontinued drug use must, in most instances, rely on prerecorded, reasonably complete documentation of use."[8]

This study used prospective ongoing records, and yet revealed systematic and substantial underrecording of infectious or flu-like illnesses, and of vitamins and over-the-counter drugs. Possible explanations for these failures to record information include:

- The woman may judge that such events as short-term febrile illnesses or use of over-the-counter drugs are too insignificant to report to a physician. This also implies that physicians do not specifically ask for this information.
- Physicians may feel that only those items known to be detrimental to pregnancy should be regularly recorded. More closed-ended questions (such as those on nausea and vaginal bleeding in this study), supplemented by open-ended ones, might improve recording practices.

Women who were still pregnant at the time of the

*Forrest JD, Singh S: Reporting of Prenatal Care in the National Natality Survey, the Vital Statistics, and the National Survey of Family Growth. Presented at the American Statistical Association Meetings, August 1985.

TABLE 2—Comparison of Reported Exposures to Medications or Short-term Illnesses Between Prenatal Women vs Postnatal Women

	Reported on Interview					
	Prenatal Women (N = 71)		Postnatal Women (N = 71)		Estimated Difference in Proportion (%)	95% CI for Difference in Proportion
	n	%	n	%		
Short-term Illness						
Upper respiratory infection	24	33.8	13	18.3	15.5	(1.3, 29.7)
"Flu" or febrile illness	13	18.3	5	7.0	11.3	(0.5, 22.1)
Nausea/Vomiting	8	11.2	1	1.4	9.8	(2.0, 17.7)
Vaginal spotting/bleeding	2	2.7	1	1.4	1.3	(−3.3, 6.1)
Other	5	7.0	5	7.0	0.0	(−8.4, 8.4)
Total reports cited*	52	14.6	25	7.0	7.6	(3.1, 12.1)
Medication type						
Prescription drugs	13	18.3	10	14.1	4.3	(−7.9, 16.3)
Over-the-counter drugs	31	43.7	22	31.0	12.7	(−3.1, 28.4)
Vitamins/supplements	60	84.5	58	81.6	3.1	(−9.5, 15.1)
Total reports cited*	104	48.8	90	31.0	6.5	(−2.9, 16.0)

*Represents total of illness classifications cited. Total denominator for short-term illness is 355 (5 complaints categories for 71 women); for medications is 213.

interview reported short-term illnesses, but not medication use, more frequently. This could reflect bias if symptomatic women were more likely to appear for prenatal care before the sixteenth week of pregnancy. However, because 94.6 per cent of our total postnatal study sample had appeared for care before 16 weeks' gestation, such a bias is unlikely. It is more likely that the postnatal women had either forgotten these events or dismissed them once a normal delivery had occurred. Either explanation argues for a potential recall bias to occur if such postnatal controls were used to compare data on early pregnancy exposures with mothers of malformed or seriously ill infants, who may be less inclined to underreport any potentially related event.[9]

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Government Report Assesses Quality of Medical Care

The Office of Technology Assessment has released an in-depth report on the quality of medical care in the US. *The Quality of Medical Care: Information for Consumers* evaluates the reliability, validity, and feasibility of specific social indicators which affect the quality of medical care that purchasers—individuals, employers, and third party payers—might use. The report assesses eight potential indicators of the quality of care:

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HIV Seroprevalence in Intravenous Drug Users: Los Angeles, California, 1986

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Abstract: To estimate the seroprevalence and investigate risk behaviors for HIV (human immunodeficiency virus) infection in intravenous drug users in Los Angeles County, a stratified, random sample of 790 clients, enrolled in either methadone maintenance or detoxification programs, were studied. Thirteen study participants (1.8 per cent) were seropositive; of these, five were homosexual/bisexual males and three were female prostitutes. Ninety-five per cent of all participants reported sharing needles. Risk reduction education efforts must be expanded before the HIV antibody prevalence rate increases. (*Am J Public Health* 1989; 79:81–82.)

Introduction

As of December 1987, 11 per cent of the 4,243 reported acquired immunodeficiency syndrome (AIDS) cases in Los Angeles County (LAC) had a history of intravenous drug use. Intravenous drug use was the sole risk behavior for 3 per cent of the cases, in contrast to the distribution of cases on the East Coast where as many as 50 per cent of AIDS patients have reported intravenous drug use as the primary risk factor.*

An estimated 120,000 active heroin addicts reside in LAC.** In 1984–85, 25,396 clients were admitted to drug treatment programs in LAC for primary parenteral substance abuse; 98 per cent were heroin users. To obtain an estimate of the prevalence of the human immunodeficiency virus (HIV) infection in intravenous drug users (IVDUs) in LAC, a serologic survey in methadone treatment programs was conducted.

Methods

Between April and July 1986, 790 study participants were selected from the approximately 25,000 IVDUs who annually receive treatment at 31 licensed outpatient methadone maintenance and detoxification centers in LAC. The sample was stratified according to clinic location and drug treatment program type (methadone maintenance or detoxification) in an attempt to obtain a representative sample of IVDUs with respect to age, race, sex, and socioeconomic status. Participants were randomly selected within stratification categories. After obtaining informed consent, participants were interviewed and a blood sample was obtained. Age, race, sex, and treatment modality were obtained for all those who refused participation. Clients who had not used drugs intravenously since 1978 were excluded.

Blood samples of participants were analyzed for the presence of antibody to HIV. Samples which were repeatedly reactive using an enzyme-linked immunosorbent assay (ELISA) were confirmed by indirect fluorescence assay (IFA) and Western blot analysis.

Data analysis was performed using Fisher's exact test, chi-square test, RxC Fisher's exact test, t-tests and exact odds ratios and confidence intervals.[1–3]

Results

Ninety-two per cent of the 790 persons offered HIV testing agreed to participate. Nonparticipants differed from the participants with respect to their drug program status; nonparticipants were more likely to be in detoxification programs than participants. Thirteen participants (1.8 per cent) had antibody to HIV as detected by ELISA, IFA, and Western blot analysis. Of the seropositives, three were female prostitutes, five were homosexual/bisexual males, one participant's spouse had died of AIDS, and one had frequented shooting galleries in New York City. The seropositive and seronegative participants did not differ with respect to age, sex, race, education, or drug treatment modality. Forty-six per cent of the seropositives had salaries greater than $20,000 per year versus 16 per cent in the seronegative group (OR = 4.45, CI = 1.21, 15.79).

Compared to heterosexual males, homosexual and bisexual males were 65-fold (CI = 3.73, 997.58) and 13-fold (CI = 1.45, 162.10), respectively, more likely to be seropositive. Twelve per cent of homosexual/bisexual males were seropositive. No seropositive males reported contact with female prostitutes. A greater proportion of seropositive females (50 per cent) were prostitutes than seronegative females (19 per cent). Four per cent of female participants who admitted to prostitution were seropositive.

The seronegative and seropositive participants began using drugs at a mean age of 21 and 20 years, respectively. Both groups injected frequently and used many types of drugs (Table 1). Approximately three-fourths of the seronegatives and two-thirds of the seropositives injected drugs at least once a day. Seropositives were more likely to speedball (inject both cocaine and heroin together) than seronegatives (OR = 2.9, CI = .82, 11.39, Fisher's p-value = .01) and to boot drugs (withdraw blood into syringe before injecting) (OR = 3.3, CI = .94, 12.93, Fisher's p-value = .03).

The majority of participants shared needles one or more times per week (Table 1). Information obtained on 37 participants who shared needles showed that two-thirds (24) shared only with a spouse, relative, or steady friend. An almost equal proportion of seropositive and seronegative participants (39 per cent and 31 per cent, respectively) injected drugs in shooting galleries. Over 63 per cent of the seronegatives and 85 per cent of the seropositives had been incarcerated at least once.

From the Los Angeles County Department of Health Services. Reprints are not available from the authors. This paper, submitted to the Journal January 22, 1988, was revised and accepted for publication June 13, 1988.

* New York City Department of Health: AIDS Surveillance Update, November 25, 1987; State of New Jersey Department of Health: AIDS Surveillance, September 1, 1987.

** Irma Strantz, DrPH, Director, Drug Abuse Program, LACHD, Personal Communication.

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TABLE 1—Drug and Substance Use in Intravenous Drug Users since 1978

Category	No. HIV Positive (%) N = 13	No. HIV Negative (%) N = 701
Drugs		
Heroin (IV)	13 (100)	694 (99)
Cocaine (IV)	12 (92)	580 (33)
Amphetamine (IV)	7 (59)	256 (37)
Barbiturates (IV)	8 (62)	470 (67)
Nitrites	2 (15)	67 (10)
Marijuana	10 (77)	480 (69)
Shooting Galleries	5 (39)	216 (31)
Sharing Needles	12 (92)	669 (95)
Booting	8 (62)	229 (33)
Speedballing	8 (62)	249 (36)
Frequency of IV Drug Usage		
≤1 time/week	1 (8)	32 (5)
1–7 times/week	3 (23)	141 (20)
7 times/week	9 (69)	518 (74)
Unknown	0 (0)	10 (1)
Frequency of Sharing Needles		
0 times/week	1 (8)	33 (5)
≤1 time/week	7 (54)	261 (37)
1–7 times/week	2 (15)	186 (27)
7 times/week	3 (23)	217 (31)
Unknown	9 (0)	4 (<1)

Discussion

The HIV seroprevalence rate of 1.8 per cent among IVDUs was surprisingly low for this large metropolitan area. Our data are consistent with studies among other IVDUs in treatment programs throughout California. Results from these studies have demonstrated seroprevalence rates ranging from 1.7 to 10 per cent.[4–6]

The epidemiologic characteristics of local AIDS cases are consistent with the low seropositivity rate observed in this study. Of the 2,026 AIDS cases reported at the time of this study, only 46 cases (2 per cent) had intravenous drug use as their only risk behavior with an additional 9 per cent of cases reported among homosexual/bisexual IVDUs. One year later in July 1987, the percentage of AIDS cases with intravenous drug use as their only risk behavior had only risen to 3 per cent. Six pediatric AIDS cases, perinatally exposed to HIV, were associated with IVDUs. The low number of cases associated with IVDUs provides additional indirect support for the low seroprevalence rate found.

Consistent with other studies, our data indicate that seropositive users are more likely to speedball and boot drugs.*** The association with booting perhaps can be explained by an increased chance of blood-virus contact in the syringe.

Our seroprevalence estimate of 1.8 per cent is a conservative one as our sample did not contain "street addicts," a group at higher risk for HIV infection. Our data may also not accurately reflect the HIV seroprevalence among minority IVDUs, especially among the Black population. Our sample also included a higher proportion of females (48 per cent) than the general population of IVDUs in treatment (35 per cent).**

The seroprevalence among IVDUs in LAC may be low because LAC does not have the large, anonymous shooting galleries that are present on the East Coast.

Our own limited data, and a recent HIV seroprevalence survey among IVDUs in Los Angeles County, showed that the vast majority (74 per cent) of users shared drugs with known acquaintances such as friends, family members, or sex partners.** Therefore, infection may be contained within these smaller groups preventing the spread of HIV to a larger population of IVDUs. LAC's large geographic metropolitan area may also isolate groups of IVDUs thereby inhibiting the spread of infection.

Although the prevalence of HIV infection in IVDUs in LAC is low, needle sharing and other high-risk behaviors in this population will likely cause this prevalence to increase in the next few years unless effective interventions are implemented.

ACKNOWLEDGMENTS

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An Outbreak of Syphilis on an Indian Reservation: Descriptive Epidemiology and Disease-Control Measures

A. Russell Gerber, MD, Leland C. King, Gerald J. Dunleavy, and Lloyd F. Novick, MD, MPH

Abstract: From January 1983 through January 1985, 88 cases of early adult syphilis and five cases of congenital syphilis were identified among residents of a southwestern Indian reservation. The male-to-female ratio of primary and secondary syphilis cases was 1.7: 1; over 90 per cent of women were of childbearing age. Using community health workers to determine the social circles of infected persons was effective in identifying persons at risk; treatment of partners of infected persons prior to development of seroreactivity contributed to the successful control of this outbreak. (*Am J Public Health* 1989; 79:83–85.)

Introduction

This report summarizes a large syphilis outbreak among residents of a southwestern Indian reservation. Conventional outbreak-control measures failed to prevent continued transmission. Identification of sexual partners by a variation of cluster techniques was vital in the control of this outbreak.

Background

Before 1983, the total number of cases for all stages of syphilis reported in any one year from this Indian reservation had never exceeded four. In 1983 and 1984, 88 cases of primary, secondary, and early latent syphilis among residents of the Indian reservation were reported to the Arizona Department of Health Services (ADHS), an increase in incidence of primary and secondary (P & S) syphilis from 32.7 cases per 100,000 population in 1982 to 163.8/100,000 in 1983 and 458.7/100,000 in 1984 (1980 estimate of the reservation population was 6,104, US Bureau of the Census). In addition, five cases of early congenital syphilis (CS) were reported.

Conventional outbreak-control measures between September 1983 and November 1984 consisted of case-finding and treatment of sexual partners identified by interview of infected persons. The report of three CS cases among reservation residents in November 1984 signaled failure to control the outbreak. This prompted an intense investigation, reevaluation of traditional control strategies, and adoption of a novel approach to secondary disease prevention using "cluster" interviewing techniques.

Clinical health care on the reservation is provided by the Indian Health Service (IHS) which maintains a hospital and a small clinic. Occasionally, reservation residents will seek primary care at off-reservation hospitals in nearby cities. Tertiary care is provided at referral hospitals in Phoenix.

From the Arizona Department of Health Services, Phoenix; and the Division of Field Services, Epidemiology Program Office, and the Division of Sexually Transmitted Diseases, Center for Prevention Services, Centers for Disease Control, Atlanta. Address reprint request to Dr. A. Russell Gerber, Division of Infectious Diseases and Immunology, Department of Pediatrics, Emory University School of Medicine, 69 Butler St, SE, Atlanta, GA 30303. Mr. King is now with the North Carolina Department of Human Resources, Raleigh; Dr. Novick is now with the Center for Community Health, New York State Department of Health, Albany. This paper, submitted to the Journal November 27, 1987, was revised and accepted for publication June 13, 1988.

Methods

The study period was January 1, 1983 to January 31, 1985. A diagnosis of P & S syphilis was made in persons with a positive serologic test for syphilis who had characteristic genital lesions or generalized rashes with palmar and plantar predilection. Early latency was diagnosed in persons without lesions who had positive serologic tests for syphilis and who had within the previous one year a nonreactive serological test, or symptoms highly suggestive of P & S syphilis, or exposure to sex partners with definitive or presumptive P & S or early latent syphilis, or a less than fourfold increase in titer on serial nontreponemal tests.[1] Suspects were defined as persons cited by patients during interview not as sex partners but as persons who might "profit from serologic testing".[2]

Difficulties in determining sexual partners through patient interview prompted screening of persons identified as high-risk through a variation of the cluster intervention technique. Community health workers were routinely interviewed regarding each case. "Close associates" were defined as persons seen frequently with infected persons by community health care workers and not named on patients' interview. All sexual partners, suspects, and close associates were informed of the availability of serologic testing for syphilis and were urged to accept treatment.

A probable case of early CS was defined as a live-born infant: 1) born to a mother with high risk for transmitting syphilis to her newborn (untreated early syphilis at the time of delivery); 2) with a positive serologic test for syphilis; plus 3) major clinical findings consistent with congenital infection.[3] A possible case of early CS was defined as above, but without the major clinical findings.

Control efforts during the outbreak included special sexually transmitted disease (STD) clinics and community outreach programs in which a total of 1,172 of the 6,104 residents of the reservation were serologically tested for syphilis, as were 80 tribal members who lived off the reservation.

Nine hundred sixty-seven persons either volunteered or were referred to the STD clinic and were serologically tested for syphilis. To increase community awareness and outreach, educational programs and serologic testing were also offered in nearby off-reservation bars, alcoholic rehabilitation centers, workplaces, and community meetings. An additional 285 native Americans were tested at these alternate locations on and off the reservation.

Results

Epidemiological Investigation

Increases in cases of P & S syphilis were first recognized during August–September 1983 when six cases were reported to ADHS (Figure 1). The majority of P & S syphilis cases occurred before May 1984. A rapid rise in syphilis cases occurred in October–December 1984. The first early CS case occurred in December 1983. Four of the five infants with early CS were born between August and December 1984. On record review, all five CS cases had deficiencies in prenatal care. Three had serologic tests for syphilis performed at the

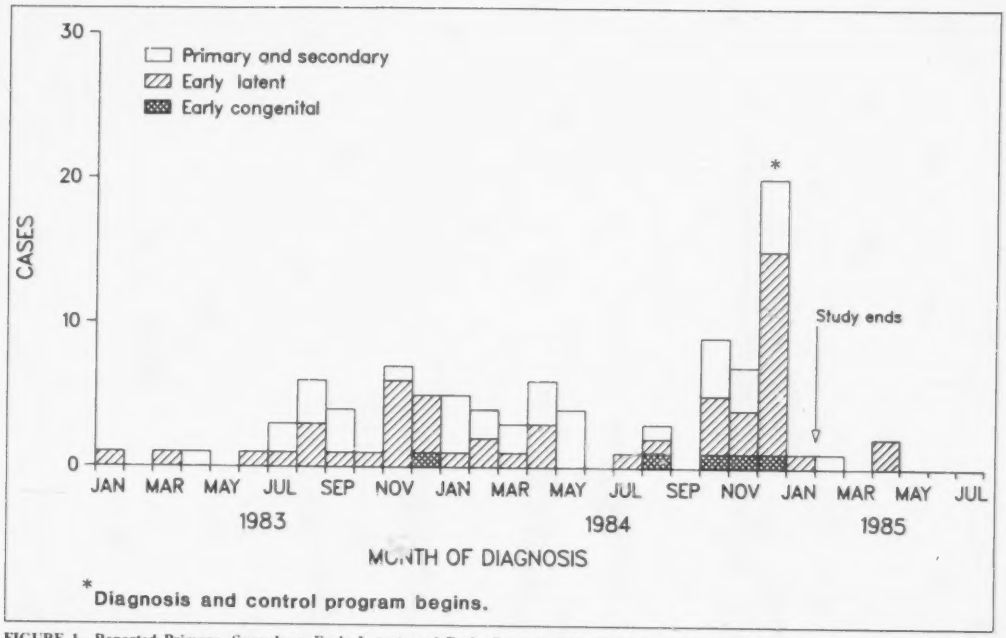

FIGURE 1—Reported Primary, Secondary, Early Latent, and Early Congenital Syphilis Cases, by Month of Diagnosis, Southwestern Indian Reservation, January 1983–July 1985

first prenatal visit, but not again during the third trimester; one did not receive prenatal care; and one with a reactive serologic test for syphilis was lost to follow-up until delivery. All infants received adequate treatment.

Of the 93 reported cases of syphilis, 38 (40.9 per cent) were P & S syphilis, 50 (53.8 per cent) were early latent syphilis, and 5 (5.4 per cent) were early CS. Of the five infants with early CS, three were classified as probable CS cases and two as possible CS cases.

Males accounted for 24 (63.2 per cent) of the persons with P & S syphilis, a male to female ratio of 1.7:1. Ages of P & S syphilis patients ranged from 14 to 48 years, with a median age of 25.5 years (males, 26.0; females, 24.0). Over 90 per cent of the females with P & S syphilis were of childbearing age (15–44 years old). The age and sex characteristics of the patients with early latent syphilis were similar to the patients with P & S syphilis.

Results of Serologic Screening

As a result of the serologic screening, a total of 24 cases were identified and treated. Three cases were identified in members of other Indian tribes living on other reservations but who had been identified as associates of residents on the index reservation; 21 cases were identified in reservation residents (Table 1). No new cases were identified in the 80 tribal members tested who lived off the reservation.

Twelve of the 20 new cases diagnosed in the STD clinic were detected among "close associates" of known cases (relative risk = 8.3; 95 per cent confidence interval: 2.9, 23.5) (Table 1). Only screening of persons actually named as partners identified new syphilis cases at a higher rate than screening "close associates."

Discussion

While the incidence rate of syphilis on American Indian

TABLE 1—Evaluation for Syphilis among Residents of a Southwestern Indian Reservation, Arizona, December 10, 1984–January 9, 1985

Groups Evaluated	Number of Patients Examined or Tested	Number of Newly Diagnosed Cases of Syphilis	Rate of Newly Diagnosed Syphilis Cases per 100 Patients Examined or Tested	RR*	(95% CL)
Sexually Transmitted Disease Clinic					
Volunteers	557	3	0.5	1.0	
Referrals Linked by Interview to Cases:					
Sexual partners	30	4	13.3	24.8	(9.1–67.4)
Suspects	110	1	0.9	1.7	(0.2–15.7)
Close Associates	270	12	4.4	8.3	(2.9–23.5)
Evaluated at Other Locations	285	1	0.4	0.7	(0.1–6.1)
Total No. Evaluated	1252	21	1.7		

*Risk of identifying a new case of syphilis relative to the rate per 100 examined or tested in a low-risk population (volunteers: 0.5/100).

reservations has decreased over the last 50 years,[4] it still remains higher than the rate for the general United States population. Preventive efforts on Indian reservations have seldom departed from the model developed for other areas—contact tracing through interview and follow-up—which relies on a case-patient naming all sexual contacts. Prompt location of persons identified by laboratory screening or interview is vital to the effectiveness of this strategy.

The control of this outbreak depended to a large degree on the utilization of knowledge possessed by community health workers to identify persons at potential risk of contracting syphilis when traditional strategies failed to prevent transmission. Close associates identified by community health workers through cluster interviewing had substantially higher rates of newly diagnosed syphilis when serologic testing was performed.

Arizona state law requires pregnant women to have prenatal serologic testing for syphilis. If prenatal care had been received and serologic tests for syphilis had been performed both at the first prenatal visit and again during the third trimester, as recommended,[3] all of the early CS cases could have been prevented.

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10,000 Homes to be Tested in Yale Radon Study

A five-year research project to help resolve some of the scientific issues relating to the radioactive gas radon and cancer will be conducted by a team of Yale University School of Medicine researchers. Funded by the National Institute of Environmental Health Sciences, located in Research Triangle Park, NC, the $3.9 million project at Yale is one of only two in the US; the other one is at the University of Utah.

This study will "carefully evaluate the serious health effects at the low end of the radon exposure scale, evaluate the interaction of radon and cigarette smoke, and consider the issue of radon as a potential risk-factor for childhood cancers," according to Stuart L. Shalat, ScD, assistant professor of medicine and epidemiology, who heads the study.

Dr. Shalat has long been interested in the effects of radiation and other issues related to occupational/environmental health. He worked in the New York State Health Department investigating illness associated with the Love Canal toxic waste site between 1978 and 1981.

Using an alpha-track monitor, the Yale researcher will measure radon levels in 10,000 houses throughout Connecticut. The device uses a film-like photo emulsion to count the number of tracks. By knowing how long the device has been measuring the presence of radon in homes, one can accurately determine the radon level. Residing in approximately one-fourth of the homes to be tested are adults with newly diagnosed cases of lung cancer, or children with new cancers. The rest of the test homes, with healthy residents, will be randomly selected. Team members will interview residents who agree to participate and will monitor their homes for a year between 1989 and 1992.

In October, the US Environmental Protection Agency identified radon, a colorless, odorless, tasteless gas, as a serious environmental health concern, and stated that radon contamination needs to be reduced in homes with levels higher than four picocuries per liter of air. A picocurie is one trillionth of a curie, a standard measure of radiation.

Radon has been a public health issue for the past five years, as more and more people have expressed concern about indoor air pollution. Because of energy-conservation measures, more people have been sealing their homes more tightly, which may increase the levels of radon and other indoor-air contaminants.

Faculty members in the departments of epidemiology and public health and internal medicine will collaborate on the study, in conjunction with the Connecticut State Tumor Registry and the Yale Comprehensive Cancer Center.

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Future APHA Meeting Dates/Sites

1989 OCTOBER 22–26 CHICAGO, ILLINOIS
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1994 OCTOBER 30–NOVEMBER 3 WASHINGTON, DC
1995 NOVEMBER 12–16 SEATTLE, WASHINGTON

NEW MASTER'S AND DOCTORAL DEGREE PROGRAMS IN HEALTH PROMOTION AND HEALTH EDUCATION LAUNCHED AT THE UNIVERSITY OF TEXAS SCHOOL OF PUBLIC HEALTH

Applications for admission in Fall 1989 to master's and doctoral programs in health promotion and health education are now being accepted.

The program of study integrates behavioral and social science theory and intervention at the individual, organizational, and community levels; program planning and evaluation; survey research and other assessment methods; health economics; and health policy. Major resources for students include the Center for Health Promotion Research and Development, the Southwest Center for Prevention Research, the Center for Prevention of Injury and Violence, the Texas Medical Center, and behavioral science programs at Rice University and the University of Houston.

For further information, write to Dr. Patricia D. Mullen, Associate Professor and Convener, Health Promotion/Health Education Module, P.O. Box 20186, School of Public Health, University of Texas Health Science Center, Houston TX 77225.

The Smoke-Free Guide:
How to Eliminate Tobacco Smoke from Your Environment by Arlene Galloway, MHSc.

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Effects of the Antismoking Campaign: An Update

KENNETH E. WARNER, PhD

Abstract: In the absence of the antismoking campaign, adult per capita cigarette consumption in 1987 would have been an estimated 79–89 per cent higher than the level actually experienced. The smoking prevalence of all birth cohorts of men and women born during this century is well below that which would have been expected in the absence of the campaign. As a consequence, in 1985 an estimated 56 million Americans were smokers; without the campaign, an estimated 91 million would have been smokers. As a result of campaign-induced decisions not to smoke, between 1964 and 1985 an estimated 789,200 Americans avoided or postponed smoking-related deaths and gained an average of 21 additional years of life expectancy each; collectively this represents more than 16 million person-years of additional life. The greatest health benefit lies in the future, however, as younger individuals reach the ages at which smoking claims its greatest toll, and as middle-aged former smokers realize relative reductions in smoking mortality risks as a result of long-term abstinence from smoking. For example, campaign-induced decisions not to smoke made prior to 1986 will result in the postponement or avoidance of an estimated 2.1 million smoking-related deaths between 1986 and the year 2000. (*Am J Public Health* 1989; 79:144–151.)

Introduction

Previous papers have presented estimates of the effect of the antismoking campaign between 1964 and 1978 on two measures of smoking behavior[1,2] and on smoking-related mortality.[3] The present paper updates the findings of the earlier studies.

Throughout this paper, the term "antismoking campaign" is used not to connote a specific set of activities oriented toward smoking prevention and cessation, but rather the collectivity of all such activities and the changing social norms that have accompanied them. In essence, "antismoking campaign" serves as a short-hand expression covering the entirety of changes in the social environment spawned by scientific and social interest in the hazards of smoking. A detailed discussion of those changes—in scientific and public knowledge, smoking attitudes and behaviors, and programs and policies—is found in the Surgeon General's 1989 report on smoking and health.[4]

Methods

Adult per Capita Consumption

An objective measure of the nation's cigarette smoking behavior, used in numerous empirical studies, is adult per capita cigarette consumption, defined as total cigarette consumption (from the Economic Research Service of the US Department of Agriculture) divided by the population over 17 years of age; its advantages and disadvantages have been discussed elsewhere.[5] To assess the impact of the antismoking campaign on cigarette consumption, as in the previous analysis[1] (and a predecessor study[5]), I employ ordinary least squares regression analysis on per capita consumption as the dependent variable. Independent variables include those conventionally found in demand analyses, such as price, and a series of variables representing major antismoking "events," such as publication of the Surgeon General's first Report on Smoking and Health[6] and the growth of the nonsmokers' rights movement. Time series data for the study cover the period 1947 through 1987. Estimated regression coefficients for the antismoking variables indicate the specific association between each "event" and per capita consumption.

From the regression results, one can also estimate the aggregate impact of the antismoking campaign in any given year by adding the values of all of the relevant antismoking variables, multiplied by their regression coefficients, to the year's actual per capita consumption. (Since the previous year's consumption is included in the regression as an independent variable, as a measure of the addiction effect, an antismoking effect is also carried forward into future years through this variable.) The resulting estimates of per capita consumption in the absence of the campaign are then compared to the actual levels of consumption to estimate the aggregate impact of the campaign in the year in question. Details on the specifics of the methodology are found in the previous analysis.[1]

Smoking Prevalence in Age- and Sex-Specific Cohorts

The second study of smoking behavior relies on *prevalence*—the most common index of the national level of smoking. In this assessment of the impact of the campaign, age- and sex-specific estimates of actual smoking prevalence derived from self-reported survey data (and therefore labeled "reported" prevalence) are compared with estimates of prevalence rates that would have been expected in the absence of the antismoking campaign (herein referred to as "expected" prevalence).

As in the earlier study,[2] the population born between 1901 and 1960 is divided into 12 age-sex cohorts, each incorporating 10 birth years. Cohort-specific smoking rates for the years 1964–78 are taken directly from the previous study, which employed retrospective self-report data* derived from the 1978 National Health Interview Survey (NHIS). For the years 1979-85, cohort-specific prevalence data are derived from the 1979, 1980, 1983, and 1985 NHIS.** For the years in which survey data are not available (1981, 1982, and 1984), linear interpolations from the available surveys are used.

To develop estimates of what the various cohorts' smoking rates would have been from 1964 through 1985, in

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*Jeffrey Harris: Personal communication.

**Office on Smoking and Health, US Department of Health and Human Services: Personal communication.

the absence of the antismoking campaign ("expected" prevalence), the analysis adopts the assumption that well-established smoking patterns or trends prior to 1964 would have persisted into the future. For men, this would mean approximate continuation of the smoking frequencies of earlier cohorts, whereas for women it would mean continuation of the growth in smoking experienced primarily since World War II.

The basic formula for developing these estimates involves adding or subtracting earlier cohorts' average percentage point changes in smoking prevalence, at comparable ages, to the 1963 base rate prevalence of the cohort in question. For example, during 1964 the 1941–50 male cohort consisted of males 14 to 23 years of age. To calculate their expected smoking prevalence in 1964, in the absence of the antismoking campaign, I add to their 1963 prevalence the average of the percentage point changes in the smoking rates of the 1931–40, 1921–30, and 1911–20 male cohorts in the years in which these cohorts were 14 to 23 years old (1954, 1944, and 1934, respectively). To calculate the 1965 expected smoking prevalence for the 1941–50 cohort, I add to the 1964 estimate the average percentage point change of the three comparison cohorts in 1955, 1945, and 1935. The same procedure is repeated for 1966 and ensuing years.

This basic procedure was applied directly to the four youngest male cohorts. For the two eldest male cohorts and all of the female cohorts, variations on the basic procedure or cohort-specific ad hoc procedures had to be applied. These are described in Appendix A.

Reported and expected prevalence rates are converted into estimates of the numbers of smokers, with and without the antismoking campaign, by multiplying cohort-specific prevalence rates by the cohort's population in a given year. For the expected number of smokers in the absence of the campaign, cohort populations are adjusted downward to reflect the differential mortality rates of smokers and nonsmokers. (The needed adjustments are developed in the "deaths postponed" analysis described below.)

Deaths Postponed by Campaign-Induced Decreases in Prevalence

The third analysis estimates the number of smoking-related deaths postponed or avoided as a consequence of campaign-induced decisions to quit smoking or not to start. In the earlier analysis, this health benefit was labeled "premature deaths avoided."[3] In the present updating, the terminology is changed to "deaths postponed" to clarify that the health benefit includes both complete avoidance and mere postponement of smoking-related deaths. Interest is in the loss of life expectancy of smokers compared to nonsmokers, whether smoking-related deaths occur at age 50 or age 70.

As in the previous analysis,[3] the number of deaths postponed is estimated by applying epidemiologic data on age- and sex-specific mortality rates of smokers, quitters, and never smokers to the cohort-specific changes in smoking prevalence discussed in the preceding section. The reduced mortality rates of quitters (compared with continuing smokers) are multiplied by the estimated number of campaign-induced quitters and non-initiators (i.e., people influenced by the campaign not to start smoking) in each cohort and in each year from 1964 through 1985. This procedure generates estimates of the number of additional smoking-related deaths that would have occurred during those years had the antismoking campaign not encouraged quitting and noninitiation. To estimate the number of life-years saved associated with each death postponed, life expectancy gained is calculated by comparing the former smoker's life expectancy with that of a continuing smoker, using life table analysis.

Much of the postponement of death due to smoking cessation and noninitiation induced by the campaign through 1985 will not be realized until later years. For example, someone who quits smoking in 1985, who would not have experienced a smoking-related death that year, might avoid a smoking-related death that would have occurred 15 years later (e.g., by avoiding the development of lung cancer). By "aging" the population of quitters and noninitiators alive in 1985 and applying to them the differential mortality rates of former smokers, one can estimate the numbers of smoking-related deaths postponed or avoided in the future as a result of quitting and noninitiation that occurred prior to the end of 1985.

Given the need to make several assumptions to develop estimates of deaths postponed, the base-case findings are subjected to four sensitivity analyses. Assumed values of each of four variables are altered in a conservative direction to determine whether the specific values selected for the base case fundamentally affect the qualitative findings of the study. The four sensitivity analyses include:

- running the estimating model with all excess smoker mortality ratios reduced by 50 per cent;
- increasing ex-smoker mortality ratios by 25 per cent,
- reducing estimates of the number of campaign-induced quitters and noninitiators by 25 per cent, and
- allowing for 50 per cent recidivism among one-year and two-year quitters and noninitiators. ("Recidivism" by a noninitiator means that someone influenced by the campaign not to start smoking in a given year decides to start smoking a year or two later.)

The precise estimating procedures for the analysis of deaths postponed from 1964 through 1985 are methodologically identical to those described in an appendix to the original study,[3] subject to one refinement. When five-year population age groupings have to be broken up to provide estimates of the population at specific ages in specific years, Karup-King coefficient formulas[7] are used; the previous analysis relied on simple averaging. Data are identical for both studies for the period through 1978. For the post-1978 period, the new age-sex cohort differences between actual and expected smoking prevalence, described in the preceding section, are used in the mortality analysis, as are more recent life table data. Other parameter values are unchanged from the original analysis.

Results

Adult per Capita Consumption

The regression model detailed in Appendix B indicates which factors reduced per capita consumption. Findings with respect to the significance of individual variables are qualitatively identical (and quantitatively very close) to those reported in the earlier study.[1] For the antismoking "events," this means that significant decreases in per capita consumption ranging from 3.5 to 7.5 per cent occurred during each year of prominent adverse smoking-and-health publicity represented by dummy variables in the regression. These include the first public smoking-and-cancer "scare" of 1953–54, publication of the Surgeon General's first Report on Smoking and Health in 1964,[6] and the three full years of the Fairness Doctrine period, 1968–70, during which broadcasters were required by the Federal Communications Commission to donate air time to the antismoking message to

"balance" pro-smoking advertising.[8] The regression also shows a strong inverse correlation between per capita consumption and growth in the number of states having laws restricting smoking in public places. Finally, the regression indicates a price elasticity of demand for cigarettes of −0.2. This means that a cigarette price increase of 10 per cent is associated with a per capita consumption decrease of 2 per cent.

The aggregate impact of the campaign on per capita consumption is shown in Figure 1, which depicts both actual adult per capita consumption and two patterns of consumption estimated in the absence of the campaign. The dashed line represents an estimate in which price changes are independent of the campaign, while the dotted line reflects the assumption that the pattern of price changes, affected most strongly by state excise taxation of cigarettes, was itself a function of legislators' reactions to the emerging evidence on smoking and health. The logic underlying the latter assumption is that, influenced by numerous state tax increases, the real price of cigarettes rose eight consecutive years following publication of the Surgeon General's first report.[1,9] A period of new-tax inactivity over the next decade has been attributed to concerns about interstate cigarette bootlegging in response to large tax-induced discrepancies in states' retail cigarette prices.[10] Consistently rising prices again since 1982 have been associated with increased antismoking sentiment in state and federal legislatures.[4]

To represent the assumption that the pattern of cigarette price was a function of the campaign, the dotted line treats price as if it had been constant (in real terms) in the years after 1963. This permits an assessment of the consumption effects of price fluctuations possibly produced as a result of the campaign. In contrast, actual experienced prices are reflected in the dashed line. The gap between the solid line (actual consumption) and dashed line (estimated consumption, actual prices) can be interpreted as a measure of the "pure publicity effect" of the campaign, while the gap between the dashed line and dotted line (estimated consumption, constant prices) measures the consumption impact of changing prices.

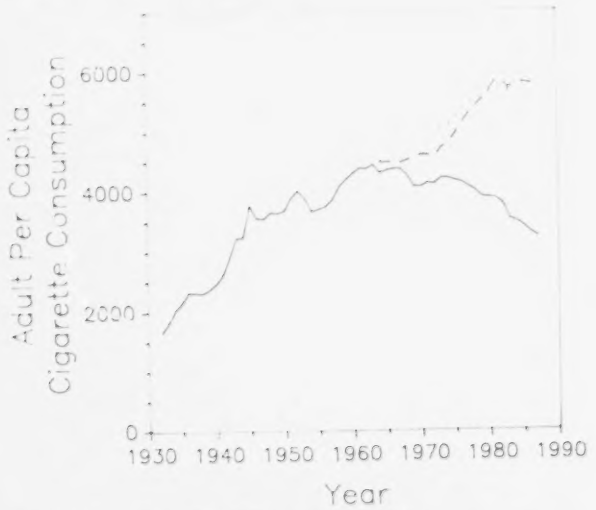

FIGURE 1—Adult per Cap. Cigarette Consumption
Solid line = actual consumption; dotted line = consumption predicted in absence of the antismoking campaign, with price held constant; dashed line = consumption predicted in absence of the campaign, with actual prices experienced

To the extent that the latter is attributable to smoking-and-health concerns of legislators, it constitutes a measure of price-related campaign-induced changes in consumption.

Each of the estimated lines exhibits a generally increasing pattern of consumption over time, in contrast to the pattern actually experienced. The analysis indicates that in 1987, per capita consumption would have been 79 to 89 per cent greater than the level actually experienced, depending on whether price changes are treated as independent of the campaign or as a function of it. The small difference suggests that, by 1987, price fluctuations contributed to decreased consumption, but to a much smaller degree than the "publicity effect."

Smoking Prevalence in Age- and Sex-Specific Cohorts

Results of the analysis of cohort smoking prevalence are presented in Tables 1 and 2 for men and women, respectively. Each cell records reported smoking prevalence for the relevant birth cohort and the relevant year, from 1964 through 1985, and the estimate of the smoking rate that would have been expected each year in the absence of the antismoking campaign.

To illustrate the nature of the findings, Figures 2 and 3 plot the numbers in Tables 1 and 2, as well as pre-1964 prevalence estimates, for two representative cohorts, one male and one female. The gap between the solid and dashed lines constitutes a measure of the impact of the campaign on smoking prevalence for these two birth cohorts.

The data in Tables 1 and 2 indicate that all 12 of the age-sex birth cohorts experienced substantial quitting, or non-initiation of smoking, as a result of the antismoking campaign. By 1985, the estimated gap between actual (reported) prevalence and that which would have been anticipated in the absence of the campaign ranged from a low of six percentage points for the eldest female cohort to a high of 28 percentage points for the youngest males. The percentage point gap is larger for each male cohort than for the same-age female cohort. For women, the percentage point gap in 1985 is larger as one moves from the older to the younger birth cohorts. For men, the gap is smallest for the oldest cohort and largest for the youngest; it is similar for the remaining four cohorts.

The "expected prevalence" figures indicate that a majority of each of the four youngest male cohorts would have been expected to be smokers in 1985 in the absence of the campaign, including two-thirds of men born 1951–60. In fact, fewer than 40 per cent of each cohort reported themselves to be smokers that year. Similarly, in the absence of the campaign, smoking prevalence in the four youngest female cohorts would have been expected to include a majority or near-majority of the women (44 per cent for the 1921–30 birth cohort to 54 per cent for the 1951–60 cohort). Yet each of these cohorts reported prevalence rates of either 27 per cent (1921–30 cohort) or 32 per cent (1931–60 cohorts).

Multiplied by their respective cohort populations, the 1985 reported prevalence figures indicate a total smoking population that year of 56 million Americans born between 1901 and 1970. In the absence of the campaign, the cohort analysis indicates that there would have been 91 million smokers. (The 1961–70 birth cohorts were added to the analysis for this purpose, with prevalence figures taken from the 1985 NHIS. "Expected" prevalence was assumed to be the same as that of the 1951–60 cohorts in 1975.)

Deaths Postponed by Campaign-Induced Decreases in Prevalence

Table 3 presents estimates of deaths postponed for each of the 12 age-sex cohorts identified in Tables 1 and 2, for the

TABLE 1—Smoking Prevalence for Males (%) 1964 through 1985, Reported[a] and Expected[a] in the Absence of the Antismoking Campaign

Birth Cohorts	Year (19--)																					Expected minus reported 1985
	64	65	66	67	68	69	70	71	72	73	74	75	76	77	78	79	80	81	82	83	84	85	
1901–10																							9
Reported	45	45	45	43	42	39	39	36	35	34	31	30	28	27	26	20	18	18	19	19	17	16	
Expected	46	46	46	45	44	42	42	40	39	38	37	37	36	36	35	34	32	31	29	28	26	25	
1911–20																							23
Reported	61	60	59	59	58	54	53	51	50	48	46	46	43	40	39	32	30	29	28	26	24	22	
Expected	62	61	60	60	59	57	56	56	55	55	54	54	53	52	51	50	49	48	47	46	45	45	
1921–30																							21
Reported	63	62	62	61	60	56	55	53	52	51	48	47	47	45	44	42	41	39	36	34	33	32	
Expected	66	66	66	65	65	64	64	64	63	63	63	62	62	61	61	59	58	57	56	55	54	53	
1931–40																							19
Reported	59	59	58	57	56	53	53	51	50	49	47	46	45	44	44	42	44	43	42	40	38	35	
Expected	62	63	63	64	64	64	64	63	63	63	62	62	62	61	61	59	58	57	56	55	55	54	
1941–50																							22
Reported	42	46	50	54	58	58	58	57	56	55	54	53	51	49	47	44	44	43	41	40	39	38	
Expected	43	48	53	57	61	61	62	62	63	63	64	64	64	64	64	63	63	62	62	62	61	60	
1951–60																							28
Reported	1	2	4	7	10	14	18	23	27	31	36	38	40	40	39	40	42	41	40	40	39	38	
Expected	2	3	5	8	11	16	21	26	32	39	44	50	54	58	61	62	64	65	65	66	66	66	

[a]See text for definition of 'reported' and 'expected' prevalence.

period 1964 through 1985. The estimated total number of deaths postponed as a result of campaign response is 789,200. Of these 112,400 occurred in 1985 alone (not shown in table). The distribution between men and women and across birth cohorts is seen in the individual cells of the table. All told, the 789,200 postponed deaths represented a savings of 16.3 million life-years, an average of 20.6 years per death postponed.

Between 1964 and 1985, Americans born between 1901 and 1960 experienced an estimated 5.7 million smoking-related deaths. In the absence of the campaign, 6.5 million deaths would have been expected. The estimated 789,200 deaths postponed or avoided represented 12 per cent of the anticipated mortality toll.

Table 4 indicates that 2.1 million deaths will be postponed between 1986 and the year 2000 as a result of campaign-induced nonsmoking that occurred through 1985. The data in the table reflect the distribution between men and women and across birth cohorts.

Aggregate results of the sensitivity analyses for the two periods (1964–85 and 1986–2000) are presented in Table 5. Although they reduce the magnitude of estimated deaths postponed, the essential qualitative finding of the study is robust: as a result of campaign-related decisions to quit smoking or not to start, many hundreds of thousands of Americans have extended their lives an average of two decades each.

Discussion

Few people question the fact that the antismoking campaign has altered smoking behavior. Adult smoking prevalence has fallen from 43 per cent at the time of the

TABLE 2—Smoking Prevalence for Females (%) 1964–1985, Reported[a] and Expected[a] in the Absence of the Antismoking Campaign

Birth Cohort	Year (19--)																					Expected minus reported 1985
	64	65	66	67	68	69	70	71	72	73	74	75	76	77	78	79	80	81	82	83	84	85	
1901–10																							6
Reported	22	21	21	20	20	19	18	17	17	16	15	15	13	13	13	13	15	13	10	8	8	8	
Expected	22	21	21	20	20	19	19	19	19	18	18	18	17	17	17	17	16	16	15	15	14	14	
1911–20																							13
Reported	36	36	36	36	35	33	33	31	31	30	29	28	27	26	26	25	26	24	22	20	19	18	
Expected	37	37	37	37	36	36	36	35	35	35	35	35	34	34	34	34	33	33	32	32	31	31	
1921–30																							17
Reported	43	43	43	42	42	40	40	39	39	39	38	38	38	37	36	31	31	30	30	29	28	27	
Expected	44	45	46	47	48	48	48	49	49	49	48	48	48	47	47	47	47	46	46	45	44	44	
1931–40																							17
Reported	44	44	44	44	44	43	43	43	42	42	42	42	41	40	39	39	35	34	34	34	33	32	
Expected	45	46	46	47	47	48	48	48	49	49	50	50	50	51	51	51	51	51	50	50	49	49	
1941–50																							20
Reported	25	30	34	38	40	40	41	41	41	41	41	40	39	38	37	35	34	34	33	33	32	32	
Expected	27	32	36	40	42	44	46	47	48	49	50	50	51	51	52	52	52	52	52	52	52	52	
1951–60																							22
Reported	1	1	2	3	6	10	13	17	22	27	31	35	38	38	37	34	32	33	34	34	33	32	
Expected	1	2	3	5	8	12	15	20	25	31	36	40	45	47	49	50	50	51	52	53	54	54	

[a]See text for definition of 'reported' and 'expected' prevalence.

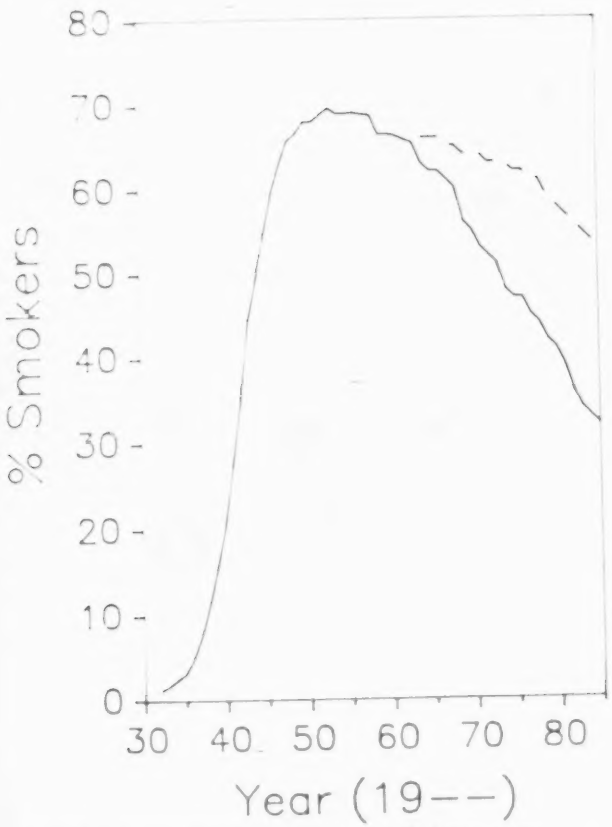

FIGURE 2—Smoking Prevalence History, 1921–30 Male Birth Cohort
Actual prevalence (solid line) and expected prevalence for 1964–85 in absence of the antismoking campaign (dashed line)

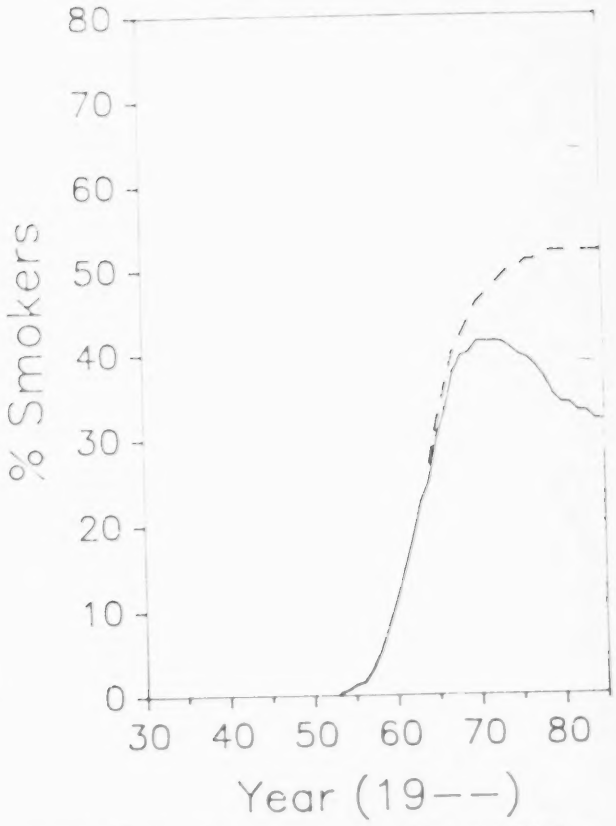

FIGURE 3—Smoking Prevalence History, 1941–50 Female Birth Cohort
Actual prevalence (solid line) and expected prevalence for 1964–85 in absence of the antismoking campaign (dashed line)

Surgeon General's first report on smoking to under 30 per cent a quarter century later; prevalence among men has almost halved. Per capita cigarette consumption has fallen annually since 1973, without exception, and is at its lowest level in 45 years; per capita tobacco consumption has reached its lowest level in a century.[4]

Yet such measures understate the impact of the antismoking campaign, since they miss the critical point that smoking prevalence and per capita consumption likely would have been rising in the absence of the campaign. For example, per capita consumption in 1987 was 27 per cent below the level that prevailed in 1963, the year prior to the Surgeon General's first report. But analysis of per capita consumption trends indicates that the 1987 value was actually less than half of what it would have been had the antismoking campaign never transpired.

This analytical perspective is most important in understanding the pattern of women's smoking. The relatively steady trend in adult female smoking prevalence from 1964 through the late 1970s, followed by relatively modest decreases in the 1980s, often has been interpreted as reflecting little response to the antismoking campaign. Yet as the cohort analysis demonstrates, in the absence of the campaign, women's smoking prevalence would have continued increasing, paralleling the diffusion pattern of smoking among men several decades earlier. As such, the stability of female smoking prevalence after the mid-1960s, at a level far below that attained by men, should be interpreted as behavioral response to the antismoking campaign.

Consistent with the observation that the campaign has affected smoking by men more than by women, the cohort analysis indicates that the percentage point gap between reported smoking prevalence and prevalence expected in the absence of the campaign is larger for each male cohort than for the same-age female cohort. Missed in this observation, however, is the comparability for men and women of the numbers of campaign-related quitters and noninitiators relative to the numbers of actual smokers. In 1985, this measure

TABLE 3—Deaths Postponed by Campaign-related Smoking Cessation and Noninitiation, 1964–85 (in thousands)

Birth year	Males	Females	Total
1901–10	103.6	16.7	120.3
1911–20	182.0	46.0	228.0
1921–30	182.7	59.6	242.3
1931–40	83.2	22.7	105.9
1941–50	44.0	15.5	59.5
1951–60	29.0	4.2	33.2
Total	624.5	164.7	789.2

of relative quit-and-nonstart rate—the ratio of the percentage point gap to actual prevalence—indicates little difference between males and females.

Growth in the gap between reported and expected smoking prevalence as one moves from the older to the younger cohorts, especially among the women, is consistent with the conventional wisdom that decisions not to start smoking are easier than decisions to quit. Decisions concerning the initiation of smoking occurred after 1964 for many members of the two youngest cohorts and well before 1984 for the older cohorts.

The peak rate of prevalence of the youngest cohort of males (42 per cent in 1980) makes this cohort the first group of men born during the 20th century never to have included a majority who were smokers. Supportive of the large gap found between this cohort's reported prevalence and the estimate of prevalence in the absence of the campaign are data on the peak prevalence of each of the older cohorts reported by Harris[11] [and unpublished data]. In the space of a single 10-year period, peak prevalence fell 16 percentage points, from 58 per cent to 42 per cent (for the 1941–50 and 1951–60 birth cohorts, respectively). In the span covered by three 10-year birth cohorts, the maximum percentage of men smoking fell by almost 30 percentage points (from 70 per cent for the 1921–30 cohort). Preliminary data from the 1987 National Health Interview Survey, combined with other NHIS data from the 1980s, suggest that the peak prevalence for the cohort of males born 1961–70 may represent a further decrease of as much as 12 or 13 percentage points, to a rate of 29 to 30 per cent.

The 70 per cent peak prevalence rates attained by the 1911–30 male cohorts suggest that the expected peak prevalence figures for males in Table 1 may be conservative. This perception is reinforced by recognition that the 1921–30 cohort achieved its peak rate of smoking in 1953, the year of the first major public smoking-and-cancer scare, and that the succeeding cohorts reached peak smoking age during periods of concern about the health consequences of smoking.

In this context, had women's smoking patterns eventually mirrored those of men approximately three decades earlier, the expected prevalence figures in Table 2 would have to be considerably larger than they are. As such, the gaps between reported and expected prevalence for women should be considered conservative measures of the impact of the antismoking campaign on women's smoking prevalence.

The deaths postponed analysis found that, by 1985, the 789,200 smoking-related deaths postponed or avoided represented 12 per cent of the smoking-related mortality that would have occurred in the absence of the antismoking campaign. In the predecessor to this study, the comparable figure through 1978 was 5 per cent.[3] The life-savings attributable to the campaign will continue to rise in the coming decades, both in absolute terms and as a fraction of total smoking-related mortality, as the cohorts most responsive to the campaign reach the ages at which smoking claims its greatest toll; in the more immediate future, campaign-related life-savings will also rise as older cohorts achieve the health benefits of long-term abstinence from smoking. This is demonstrated by the estimate of 2.1 million deaths postponed between 1986 and the year 2000 among those who had quit smoking or not started through 1985. The benefits of the campaign through 1985 alone substantially understate the long-term health dividend of campaign-induced behavioral changes that have already taken place.

TABLE 4—Deaths Postponed 1986–2000 by Campaign-related Smoking Cessation and Noninitiation through 1985 (in thousands)

Birth year	Males	Females	Total
1901–10	18.8	5.9	24.7
1911–20	194.6	83.8	278.4
1921–30	429.5	162.6	592.1
1931–40	315.3	128.6	443.9
1941–50	290.5	109.8	400.3
1951–60	241.7	76.3	318.0
Total	1490.4	567.0	2057.4

TABLE 5—Sensitivity Analyses

Sensitivity Test	Deaths Postponed (thousands) 1964–1985	1986–2000
1 Reduce excess smoker mortality ratios by 50%	519.8	1,416.1
2 Increase ex-smoker mortality ratios by 25%	581.1	1,876.5
3 Reduce campaign-induced quitters and noninitiators by 25%	592.8	1,547.7
4 Allow 50% recidivism among 1- and 2-year quitters and noninitiators	406.0	1,044.8

The analysis concluded that men have realized a much greater collective health benefit from their responses to the antismoking campaign than have women. This finding reflects more men than women having quit and not initiated smoking as a result of the campaign, and women having lower death rates and smoker mortality ratios than men. As women's smoking behaviors have approached those of men in recent years, the disparity between male and female smoker mortality ratios has diminished.[4] As such, this study's conservative assumption about women's smoker mortality ratios may have produced an underestimate of the health benefit of the campaign for women. Particularly in the younger cohorts, if women respond to the relatively new emphasis on the health consequences of their smoking,[12] the male-female differential of premature deaths avoided would be expected to decrease in the future.

The analysis of deaths postponed presents an incomplete picture of the health consequences of the antismoking campaign because it ignores the smoking-related morbidity and disability avoided. Moreover, it does not address directly one of the major behavioral impacts of the campaign: the shift from the unfiltered cigarette of the 1940s and '50s to filtered cigarettes and, more recently, to the low tar and nicotine (t/n) product.[4]

On the surface, this profound change might appear to represent a positive contribution to diminishing the toll of smoking. Yet recent evidence indicates that smokers compensate for reduced nicotine yields, often unconsciously, by such methods as smoking more cigarettes, inhaling more deeply, taking more puffs per cigarette, and blocking air ventilation holes in the filters.[13] Low-yield cigarette smokers are also exposed to additives that may introduce new health risks.[14] Further, the availability of filters and low-yield cigarettes may have encouraged many smokers to continue smoking when they otherwise would have quit. It is also possible that they have encouraged certain groups to start smoking in larger numbers, perhaps most notably young

females.[15] Each of these possibilities implies increases in smoking-related morbidity and mortality relative to that which would have occurred in their absence. As such, the net health consequences of changes in the cigarette product are unknown.

The most obvious concern about smoking is the continuing population of more than 50 million smokers. While, in the absence of the campaign, the number of smokers today would exceed 90 million, the remaining population of smokers assures that the enormous burden of smoking will persist into the 21st century. The fact that tens of millions continue to smoke, when the vast majority of them wish that they did not, illustrates the tenacious grip that nicotine places on its users.[14] This sobering reality places the impressive successes of the antismoking campaign in context, but it does not diminish them.

APPENDIX A

Cohort-Specific Procedures to Estimate Smoking Prevalence in the Absence of the Antismoking Campaign

The text explained the basic procedure for estimating cohorts' smoking prevalence rates in the absence of the antismoking campaign. For the two oldest male cohorts and for each of the female cohorts, deviations from the basic procedure were required. Additional minor adjustments were also required for the most recent years of the study.

For the more recent years of the study, comparison cohorts often reached comparable ages *during* the antismoking campaign. For example, the three comparison groups for the 1951–60 male cohort were the 1911–40 cohorts. To use the 1931–40 cohort in estimating the 1951–60 cohort's expected prevalence in 1984 and 1985 would place the former at comparable age in 1964 and 1965, the first two years of the campaign. Thus, for these two years, the prevalence changes of the two older cohorts alone were employed. In general, whenever a comparison cohort was of comparable age during the campaign, it was dropped from the analysis and the remaining one or two cohorts' prevalence changes were used.

When adequate comparison groups were not available, cohort-specific ad hoc procedures were adopted. In the instance of the 1921–40 male cohorts, this was needed only to estimate expected prevalence in 1984 and 1985. In both cases, the average decrease of the three preceding years was used.

For the 1911–20 male cohort, the basic procedure was followed through 1978 (using only the 1901–10 cohort as the comparison group, and making minor adjustments to compensate for the problem of understatement bias discussed by Harris[11]). Thereafter, expected prevalence was decreased 0.924 percentage point per year, the average of the annual declines from 1950–69 for the 1901–10 cohort.

The 1901–10 male and female cohorts had no useful comparison groups, since retrospective reporting of smoking histories in 1978 by people born before 1901 produces a significant underestimation bias[11]; in addition, the smoking histories of these earlier generations were not "models" for successor cohorts. Thus, for 1964–78, the expected prevalence rates were estimated by taking an arbitrary reduction of 50 per cent in the reported percentage point decrease in the rate of smoking. For the men, for expected prevalence after 1978, no additional differential change was assumed; i.e., an annualized rate of decline was selected to make the gap between expected and reported prevalence the same in 1985 as it was in 1978. For the women, a three-point decline between 1978 and 1985 was assumed, as it was for the 1911–30 female cohorts during these years. This decline was more gradual than for the same-age men, but the men's prevalence figures (both reported and expected) were higher than the women's in all cases.

For the earlier years (1964–78), there was no useful comparison group for the 1911–20 female cohort, since their smoking behavior differed radically from that of their predecessors, among whom smoking was quite rare. Arbitrarily, the cohort's 1956–63 smoking rate (37 per cent) was continued through 1967 and then decreased at a decelerated rate compared to the decline in reported smoking prevalence.

For the 1921–30 female cohort, given the emergence of the women's liberation movement, with its smoking manifestations, we assumed a continuing gradual increase in smoking prevalence through 1971, at which point a peak rate of smoking 49 per cent was assigned. For the remaining years, a very gradual rate of decline in expected prevalence was assumed.

For the three youngest female cohorts (born 1931–60), the basic estimation method was employed for the years 1964–78, subject to a slight modification: we assumed that the intercohort upward trend in peak smoking prevalence would have led these cohorts, in the absence of the antismoking campaign, to reach maximum smoking rates slightly above 50 per cent. This seemed a conservative assumption, given that actual peak prevalence for males of earlier cohorts exceeded 60 (and even 70) per cent.

For the years after 1978, the basic method did not apply, for lack of adequate comparison groups. Each of these three female cohorts was treated differently, reflecting their age differences. For the eldest (1931–40), expected prevalence was decreased by a total of two percentage points. Same-age males' expected prevalence fell by seven points but remained higher (at 54 per cent) in 1985 than that of the females (49 per cent). The more gradual decline among women seemed appropriate in light of the later diffusion of smoking among women and the later age of peak prevalence for this cohort.

For the 1941–50 female cohort, expected prevalence was held constant at 52 per cent during 1979–85. For the same-age males, expected prevalence fell slightly (64 to 60 per cent). Again, the later diffusion of smoking among women and the later age of actual peak prevalence made this assumption seem reasonable and conservative, since expected prevalence generally remained about 10 points below that of the same-age males.

For the youngest group of women (1951–60), expected prevalence growth of five points from 1978 through 1985 was assumed, equaling the expected prevalence growth for the same-age men and taking expected prevalence to 54 per cent, fully 12 percentage points below that of the same-age men.

Since, of necessity, many of these patterns of expected prevalence involved arbitrary judgments, their overall reasonableness was assessed by aggregating them and comparing them with the estimates of adult per capita consumption described in the text. This procedure suggested that the assumed patterns are reasonable in their representation of smoking prevalence trends in the aggregate. Obviously, the procedure cannot be used to determine whether individual cohorts' patterns are reasonable.

The major use of this cohort analysis, in evaluation of the number of deaths postponed as a result of campaign-response, includes a sensitivity analysis in which the gaps between actual and expected prevalence are diminished by 25 per cent. As such, the principal qualitative conclusions drawn

from the analysis of prevalence effects of the campaign are unlikely to be affected by possible misjudgments in evaluation of expected prevalence rates.

A few additional details on the estimating procedures, including explanations of why they are believed to be conservative, are presented in the appendix to the predecessor study[2] and in a document available from the author on request.

APPENDIX B

Adult per Capita Consumption Regression

The per capita consumption regression results are as follows:

$$C_t = -3565.2 - 7.83^{*}P_t + 0.74^{*}C_{t-1} + 1390.7^{*}Y_t - 153.66^{**}D53_t - 279.37^{*}D54_t - 181.36^{*}D64_t - 158.58^{**}D68_t - 278.17^{*}D69_t - 140.63^{**}D70_t - 23.92^{*}NSTATES_t \quad R^2 = .976$$

[* $p < .01$; ** $p < .05$]

where —C_t is adult per capita cigarette consumption in year t (mean = 3944)

—C_{t-1} is lagged consumption, which captures the effect of addiction

—P_t is relative real cigarette price in year t (index with $P_{1967} = 100$; mean = 92.32)

—Y_t is the natural logarithm of the last two digits of year t and reflects increases in the smoking population due to the diffusion of the behavior, particularly among women, and increases in smokers' daily consumption levels (mean = 4.18)

—$D53_t$, etc., are dummy variables, for the years indicated by the digits (except for $D54_t$), which provide estimates of the impact on consumption during these years of major antismoking activity (each = 1 in the year indicated by the digits, 0 otherwise)

—$D54_t$ is 0 before 1954 and $0.5^{(t-1954)}$ in 1954 and thereafter, a dummy which equals 1 in 1954 and has a continuing, though rapidly diminishing, additional effect that reflects additional publicity through the mid-1950s

—$NSTATES_t$ is the number of states having smoking restriction laws in effect in year t, a proxy measure for the effectiveness of the nonsmokers' rights movement (mean = 20.19)

The variables and data sources are identical to those used in the previous study,[1] except that the proxy measure of the success of the nonsmokers' rights movement in the earlier study was the percentage of the adult population residing in states with laws restricting smoking in public places, calculated by the author. The new variable was preferred because it comes from an independent source[16] (updated by the Office on Smoking and Health).

As noted in the text, findings with respect to the significance of individual variables are qualitatively identical to those reported in the previous study. The relatively small differences in the estimated coefficients in the two studies primarily reflect the addition of nine years to the period studied. The most striking difference reported in the Results section is the apparent reduction (in absolute value) in the price elasticity of demand from −0.37 to −0.2. While this might reflect a real change in the elasticity, representing a smoking population that is older and more addicted, it could also simply reflect the vagaries of time series analysis. Both figures are within the range of price elasticities found in other studies.[4]

The Results section also reports the strong inverse correlation between consumption and growth in the number of states with laws restricting smoking in public places. As observed in the previous study, this correlation should not be interpreted as implying that passage of nonsmokers' rights laws leads to reductions in cigarette consumption.[1] In fact, a new econometric study has concluded that the direction of causality might be the opposite: that is, decreases in consumption may be causing the growth in nonsmokers' rights legislation.[17]

Readers interested in further details on the interpretation of the regression results should consult the previous paper.[1]

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International Comparisons of Trends in Cigarette Smoking Prevalence

JOHN P. PIERCE, PhD

Abstract: Data on smoking prevalence since 1974 are presented for the United States, Canada, Great Britain, Australia, Norway and Sweden. During this period, sex-specific prevalence has decreased in all the countries studied, with the exception of Norway, where women showed an increase. There was also a considerable decline in uptake of smoking by the young over this period, suggesting that the observed decline in prevalence is likely to continue. In the United States, the rate of decline in adult smoking prevalence has been linear. This linear pattern is probably similar in prevalence in most other countries studied, with the notable exception of Australia, which demonstrated no change for the majority of the period.

Among the six countries studied, the United States had neither the lowest smoking prevalence nor the fastest rate of decline over the period. Differential patterns of change infer that the successful public health interventions in some countries are not being applied in others. While the lack of change in Australia prior to 1983 is surprising, this was followed by a sizable drop in smoking prevalence for both higher and lower educational groups in conjunction with the introduction of mass media-led antismoking campaigns. Most of the other countries report an ever increasing gap in prevalence between higher and lower educational groups. These findings suggest that all countries might benefit from a greater exchange of antismoking ideas and public health action. (*Am J Public Health* 1989; 79: 152–157.)

Introduction

It is now over 25 years since the first major authoritative reviews of the serious health consequences of smoking[1,2] appeared in the public health literature. Since that time, the degree of public health mobilization and the type of activity have differed considerably across countries in the developed world. The most recent Surgeon General's report reviews progress in the United States over the last 25 years.[3] Of interest is how the United States has performed in comparison to other nations in achieving a reduction in the prevalence of smoking.

This paper reviews recent trends in reported cigarette smoking prevalence in six developed countries which have used different mixes of public health action to reduce cigarette smoking in their societies. Smoking trends across countries can be compared to identify those that have exhibited significantly higher rates of change or sudden shifts in prevalence. Either of these phenomena suggest the presence of a more effective intervention. If such occurs, then an analysis of public health action in that country could lead to suggestions for improvement in other countries.

In comparing cigarette smoking in differing countries, care needs to be taken to ensure that major sociodemographic differences that are associated with smoking do not confound the comparison. For the United States, the two most important differences are gender and education.[4,5] Accordingly, in this paper, I report smoking prevalence by each of these variables. Furthermore, as the proportion of each population taking up the habit is a crucial variable affecting future smoking prevalence,[6] I report a measure of sex-specific uptake for each country.

Methods

The major markers of behavioral change in smoking are the sex-specific adult prevalence, uptake of smoking by the young, and trends in smoking behavior in different educational categories. The goal measure of adult prevalence in this paper is the prevalence of smoking in those aged 20 years and over. As in the trend analyses for the United States,[3,4] uptake of smoking is indicated by the proxy variable "prevalence in age 20 to 24 years." The advantage of this measure is that differences due to variations in wording (such as daily smoking or any smoking) should be minimized; at least in the United States, 90 per cent of smokers report starting to smoke regularly before age 21.[3] This proxy for uptake will give a lagged estimate of what is actually happening by as much as five years.[4]

Details of data collection procedures and definitions for these two variables are presented below for each of the six countries. As educational level and education categorization of citizens in different countries differ considerably, this variable is presented separately to allow cross-cultural comparisons of categorizations.

United States

In the United States, the National Center for Health Statistics undertakes a regular household survey on the health practices of the nation. Since 1974, there have been regular smoking supplements to these surveys which report self-reported smoking behavior on a representative random sample of respondents. Smoking status was defined from two questions: "Have you smoked at least 100 cigarettes in your entire life?" and "Do you smoke cigarettes now?"[7] In the years previous to 1974, similar smoking data were obtained, often from another (proxy) adult living in the same household; this methodological difference is the reason for starting these trends in 1974. The response rate for the tobacco supplement has been reported as varying between 87 and 90 per cent.[4] Comparison of estimates of consumption from self-reported smoking with estimates from tax data[8] demonstrates a consistent 30 percentage point difference between the two estimates. Biochemical testing in community studies[9,10] suggests that this difference is not a result of people misreporting their smoking status.

Great Britain[11]

The Social Survey Division of the Office of Population Censuses and Surveys in Great Britain has conducted an annual General Household Survey since 1974. Sample size for these surveys has been between 12,000 and 14,000 households. Interviews are sought with all members of the household over 16 years of age and, where this is not possible, proxy interviews are accepted from a near relative.

From the Office on Smoking and Health, Center for Chronic Disease Prevention, Centers for Disease Control, US Public Health Service. Address reprint requests to John P. Pierce, PhD, Chief, Epidemiology Branch, Office of Smoking and Health, 5600 Fishers Lane, Rockville, MD 20857. This paper, invited by the editor, was accepted for publication October 25, 1988.

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This has occurred in approximately 5 per cent of households. Responses have been obtained from 81 per cent to 84 per cent of households in each survey year. Questions on smoking habits were included on each survey from 1972 to 1976 and every second year thereafter. Cigarette smoking status is defined from responses to the following questions: "Do you smoke cigarettes at all nowadays?" "Have you ever smoked cigarettes regularly?"

Other surveys have been commissioned regularly by the Tobacco Advisory Council and the Office of Population Censuses and Surveys. The first set of surveys are undertaken annually on quota samples of approximately 10,000 people. These data are adjusted to overcome differences in estimates between the self-reported and the tax data.

Canada[12]

Statistics Canada is responsible for an annual Labor Force Survey which covers the civilian non-institutionalized population 15 years of age and over from the 10 provinces of Canada. In the survey, 56,000 dwellings are sampled each month with sub-samples being interviewed each month for six months in a panel design. The Smoking Behavior of Canadians Survey has been conducted as a supplement to the Labor Force Survey; the sample was approximately 31,000 individuals in 1986 with a reported response rate of 95 per cent. Proxy respondents are accepted and smoking status of individuals is decided from responses to the following questions: "Has (the chosen respondent) ever smoked cigarettes, cigars or pipe?" "At the present time, does (the chosen respondent) smoke cigarettes?" "At the present time, does (the chosen respondent) smoke cigarettes regularly?" In 1986, proxy respondents made up 30 per cent of the sample.

Australia[13,14]

Surveys of tobacco smoking in Australia have been commissioned five times since 1974 by the Anti-Cancer Council of Victoria using the national commercial pollster, Roy Morgan Research.[9] Cluster samples of 10 interviews were obtained by approaching consecutive households starting from a stratified random sample of houses drawn from electoral rolls. Data were collected from respondents ages 14 years and over. These surveys report a 40 per cent non-response rate, half of which comes from those we were unable to contact and the other half from refusals to participate in the well-known weekly "omnibus" survey on consumer products. The representativeness of the sample was demonstrated by comparing demographics obtained for subdivisions with those obtained from published Census data.[10] Validation of self-reported smoking was shown by comparison with saliva cotinine on large sub-samples at two points in time.[4]

Respondents were shown a card and asked to classify themselves into one of 10 smoking categories made up from current smoking of the different combinations of cigarette, cigar, and pipe smoking. Smoking prevalence used in this paper is derived from the following categories "Smoke only cigarettes" "Smoke cigarettes and also cigars/pipe."

Scandinavian Countries

Each year, in each of the Scandinavian countries, surveys are conducted of nationwide, representative population samples between the ages of 18 and 70 years drawn from a population registry. In Sweden,[15] in-person household interviews on samples of approximately 2,000 people were conducted from 1976 to 1983 with mailed questionnaires being used from 1984 to 1987. Between 1976 and 1983, response rates varied from between 80 and 82 per cent. Since 1984, a direct random sample of equivalent size drawn from the nationwide registry received a mailed questionnaire with a letter emphasizing the importance of self-completion and guaranteeing anonymity. Non-respondents are given two mailed reminders after which a random 50 per cent are interviewed by telephone. The estimated weighted response rate has varied between 85 and 90 per cent. Each of these countries uses the questions recommended by the World Health Organization to categorize smokers: "Have you ever smoked?"; "Have you ever smoked daily for six months or more?"; "Do you now smoke daily, occasionally or not at all?"; "What kinds of smoking material do you use (cigarettes, other)?"

The Norwegian Central Bureau of Statistics conducts annual in-person household interviews on a representative sample (approximately 2,500 each year) of the Norwegian population between 16 and 74 years of age.[16] Smoking questions are the same as those used for Sweden.

Educational Socioeconomic Categorization

Educational level of smokers is reported for each country except the United Kingdom which prefers to report occupationally based categories of social class or major socioeconomic group. The Appendix presents approximate equivalent categories for each country. In all countries there is a natural division of education between secondary or high school level and post-secondary. However, the division of post-secondary into college graduate and less than college graduate has varied over the years in differing countries. Similarly, the actual level of certification for satisfactory completion of high school education has varied within a given country over time. Accordingly, caution should be used in interpreting these data. The data are satisfactory for assessing whether there is a smoking differential across educational categories and for identifying whether there has been a differential trend across such categories as has been suggested in the United States.[4]

Analysis

For the United States data, prevalence estimates from 1974 to 1985 were weighted to reflect the US population and age standardized to the 1985 age distribution. Standard errors for prevalence estimates were adjusted for sampling design by using variance curves provided by the National Center for Health Statistics.[4] The change in these data was equally well fit by both a linear and a logit model[6]; for simplicity, the linear model is preferred. Weighted least squared regression analysis was used to calculate the average yearly change in prevalence over the period. The effect of age adjustment and using weighted analysis altered the estimated yearly change by less than 0.05 per cent per year.

For other countries, the reported prevalence has been weighted to reflect the population. Neither age standardization nor weighted analyses were possible from the available data. Using the linear pattern in the United States as justification, a simple linear model was fitted to the data from each country to arrive at an estimated change per year. Without the relevant standard errors and with age adjustment of the population for different years, this analysis must be considered to be crude and should only be used to identify differences which are orders of magnitude away from those reported for the United States.

Results

Smoking Prevalence

By 1974, smoking prevalence in Adult males in the United States, Australia, and Canada was in the low 40s (Table 1). These levels were considerably lower than those for Great Britain and Norway where over half of the adult male population was smoking. However, by the mid 1980s, with the exception of Norway, all six countries reported male smoking prevalence in the low 30s. For the United States, a linear model fits the 1974 to 1985 data extremely well (R^2 = 0.97) and smoking prevalence has been decreasing each year by 0.91 of a percentage point (standard error 0.03). According to this model, the 1987 estimate is 31.7 which is exactly the same as the observed estimate. There was considerable variation in the average rate of change in smoking prevalence calculated for the other countries. The estimated change was largest in Great Britain at 1.35 percentage points per year. Canada, Norway, and Sweden report changes of the same order of magnitude as the United States. Among Australian males, the linear model indicates that there has been little change in smoking prevalence. The observed data suggest that minimal change occurred between 1974 and 1983. However, there was a drop of six percentage points between 1983 and 1986, suggesting that the linear model is inappropriate to describe smoking trends among Australian males over this period.

In contrast to males, female smoking in 1974 had not risen to much over 30 per cent of the population in any country except Great Britain (Table 2). By the mid 1980s, smoking prevalence had declined in adult women in all countries except Norway and perhaps Australia. In the United States, the linear model fitted to the 1974 to 1985 observed data explained 88 per cent of the variation and the change per year was 0.33 per cent (SE 0.06). According to the linear model, the 1987 estimate should have been 26.9 per cent, 0.1 per cent higher than that observed. This rate of change is lower than that estimated for Great Britain, Sweden, and Canada. While the rate of change in almost all countries is well below that reported for males, by the mid 1980s in all countries except Sweden, there were still fewer female than male smokers.

TABLE 1—Trends in Smoking Prevalence for Males, ages 20 and Over

Year	United States	Great Britain	Australia	Canada	Norway††	Sweden†
1974	43.4	52	42.3		54.0	
1975				45.6	49.6	
1976	42.1	46	40.9		49.9	36
1977	40.9			43.0	44.9	32
1978	39.0	45			46.6	33
1979	38.4			40.4	45.2	31
1980	38.5	43	43.3		43.7	26
1981				38.7	40.5	27
1982		39			41.2	26
1983	35.5		38.9	35.7	44.6	26
1984		37			43.3	27
1985	33.0				40.3	25
1986		35	32.9	32.3	43.8	24
1987	31.7*				41.3	24
Change per year	−0.91**	−1.4‡	−0.05*†	−1.21‡	−0.77‡	−0.99‡

†18–70 years only, includes only daily smokers for 6 months or more
††20–70 years only
*provisional data only
**obtained from weighted regression analysis of age standardized data from 1974 to 1985[6]
‡obtained from simple linear regression of yearly prevalence data
SOURCES: United States: Fiore, *et al.*[4]; Great Britain: Wald, *et al.*[12] and Ebi-Kryston personal communication; Australia: Hill,[13] Pierce,[14] and personal communication; Canada: Millar,[12] and personal communication; Norway: Lund, personal communication; Sweden: Ramstrom[15] and personal communication

TABLE 2—Trends in Cigarette Smoking Prevalence for Females, 20 Years of Age and Over

Year	United States	Great Britain	Australia	Canada	Norway††	Sweden†
1974	31.4	41	31.9		31.4	
1975				32.0	32.1	
1976	31.3	38	34.5		31.2	34
1977	31.4			31.8	29.7	31
1978	29.6	37			31.5	34
1979	29.2			30.7	33.1	32
1980	29.0	37	32.9		30.5	26
1981				29.6	31.0	31
1982		34			34.8	30
1983	29.4*		31.4	29.3	32.3	28
1984		32			35.5	30
1985	28.0				32.5	31
1986		31	30.6	26.6	32.4	30
1987	26.8*				33.3	27
Change per year	−0.33**	−0.8‡	−0.02‡	−0.49‡	+0.20‡	−0.58‡

†18–70 years only, includes only daily smokers for 6 months or more
††20–70 years only
*provisional estimate only
**obtained from weighted regression analysis of age standardized data 1974–1985[6]
‡obtained from simple linear regression of yearly prevalence data
SOURCES: United States: Fiore, *et al.*[4]; Great Britain: Wald, *et al.*[12] and Ebi-Kryston personal communication; Australia: Hill,[13] Pierce,[14] and personal communication; Canada: Millar,[12] and personal communication; Norway: Lund, personal communication; Sweden: Ramstrom,[15] and personal communication

Uptake of Smoking

With the exception of the United States and Sweden, the prevalence of smoking among males from 20 to 24 years of age approximated 50 per cent in the mid-1970s (Table 3). During the mid-1980s, however, there was a major decline in the uptake of smoking in the United States, Australia, and Canada. In the United States, the linear model explained 75 per cent of the variation in the observed 1974 to 1985 data. The rate of decrease was 1.03 percentage points per year (SE 0.27). According to this model, the 1987 figure should be 31.3 which is slightly higher than the observed 31.1 per cent. Considerable variability was again evidenced in the Australian data with a high 56.5 estimate for 1980 which meant that the slope of the regression line was close to zero.

Prevalence of smoking in young females (Table 4) varied considerably between countries in the mid-1970s with the highest rate of smoking occurring in Scandanavia. Thereafter, higher estimates of smoking prevalence were recorded in almost all countries for at least one year before prevalence started to decrease. In the United States and Australia, the mean change over the period was positive. In the case of the United States, the linear model had a rate of change of +0.11 percentage points per year (SE 0.27). Smoking prevalence in 1987 was 28.1 which is substantially lower than the estimate from the linear model. In all countries except Australia, the mid-1980s estimate was below that of 1974. In nearly all countries, once the prevalence began to drop it did so at a rate equal to or greater than that of young males over the whole period. This was particularly evident in the Canadian data.

TABLE 3—Trends in Smoking Prevalence for Males, 20 to 24 Years of Age

Year	United States	Great Britain	Australia	Canada	Norway	Sweden†
1974	44.3	52	49.2		52.5	
1975				48.3	41.8	
1976	45.9	47	46.4		49.2	39
1977	40.4			40.7	37.6	28
1978	38.5	45			47.8	36
1979	37.7			39.8	41.0	27
1980	40.0	44	56.5		44.2	25
1981				40.8	36.2	20
1982		41			41.4	17
1983	36.9		42.5	37.3	37.8	17
1984		40			39.1	25
1985	31.0				43.5	14
1986		41	39.8	32.0	36.5	25
1987	31.1*				36.4	14
Change per year	−1.03**	−0.9‡	−0.06‡	−1.23‡	−0.80‡	−1.71‡

†18–24 years, includes only daily smokers for 6 months or more
*provisional estimate only
**obtained from weighted regression analysis of age standardized data from 1974 to 1985[6]
‡obtained from simple linear regression of yearly prevalence data
SOURCES: United States: Fiore, *et al.*[4]; Great Britain: Wald, *et al.*[12] and Edi-Kryston personal communication; Australia: Hill,[13] Pierce,[14] and personal communication; Canada: Millar[12] and personal communication; Norway: Lund, personal communication; Sweden: Ramstrom[15] and personal communication

TABLE 4—Trends in Cigarette Smoking Prevalence for Females, 20–24 Years of Age

Year	United States	Great Britain	Australia	Canada	Norway	Sweden†
1974	35.4	44	37.6		45.4	
1975				38.3	45.4	
1976	34.2	45	43.2		42.2	49
1977	37.4			45.2	37.9	44
1978	32.5	43			45.4	44
1979	34.0			42.3	47.0	36
1980	32.5	40	39.7		42.0	27
1981				39.9	36.2	40
1982		40			49.4	34
1983	37.0		43.3	37.3	37.8	29
1984		36			43.1	28
1985	32.5				43.5	33
1986		38	40.6	31.3	34.2	42
1987	28.1*				33.1	27
Change per year	−0.11**	−0.7‡	−0.23‡	−1.39‡	−0.88‡	−1.31‡

†18–24 years, includes only daily smokers for 6 months or more
*provisional estimate only
**obtained from weighted regression analysis of age standardized data from 1974 to 1985[6]
‡obtained from simple linear regression of yearly prevalence data
SOURCES: United States: Fiore, *et al.*[4]; Great Britain: Wald, *et al.*[12] and Edi-Kryston personal communication; Australia: Hill,[13] Pierce,[14] and personal communication; Canada: Millar[12] and personal communication; Norway: Lund, personal communication; Sweden: Ramstrom[15] and personal communication

Smoking by Educational/Socioeconomic Level

In each of the countries studied which reported smoking data by educational level, the prevalence of smoking decreased with increasing education (Table 5). Furthermore, there was approximately a twofold difference in smoking prevalence between the highest and lowest educational categories.

TABLE 5—Trends in Cigarette Smoking Prevalence by Education Level, 20 Years of Age and Over

Educational Level	United States	Australia	Canada	Norway††	Sweden†
High					
Group 1	16.3	19.9	19.9		
Group 2	26.1	29.8	26.0	22	25.8
Medium	33.1	32.9	34.2	37	29.0
Low	35.7	36.3	31.6	43.5	34.1

†18–70 years only, includes only daily smokers for 6 months or more
††20–70 years only
SOURCES: United States: Pierce, *et al.*[5]; Australia: Hill, personal communication; Canada: Millar[16]; Norway: Lund, personal communication; Sweden: Ramstrom, personal communication

A similar trend was observed across socioeconomic groups in the United Kingdom. In 1984, smoking prevalence in professional men and women was 17 per cent and 15 per cent, respectively. For males, smoking prevalence rose inversely with socioeconomic category to a high of 49 per cent for the unskilled, manual category. Smoking prevalence in women was in the high 20s for non-professional white-collar workers and in the high 30s for all workers in lower socioeconomic categories.

Trends in Smoking by Educational Level

Trend information for the highest and lowest categories of education are presented in Table 6 for four countries. In the United States, the linear model for college graduates explained 92 per cent of the variation in the observed data. The estimated reduction in smoking prevalence per year is 0.91 percentage points (SE 0.13). This model predicted that the 1987 prevalence would be 16.5 per cent which was slightly higher than the observed 16.3 per cent. For those with less than high school education, the linear model accounted for 85 per cent of variation in the data. The annual rate of decline was 0.19 percentage points (SE 0.03). This model predicted that smoking prevalence in this group would be 33.8 percent in 1987, considerably lower than the observed estimate of 35.7 per cent. Thus, the difference in the rate of reduction between the highest and lowest education levels was at least ninefold. A twofold difference was reported in Canada and a threefold difference in Norway. Australia was the only country in which smoking prevalence decreased approximately equally in both the highest and lowest categories.

Discussion

Trends in smoking prevalence have been compared for six countries known for their public health action to reduce smoking in their societies. In each country, there was a marked difference in the rate of change between males and females as the historical gap between the sexes diminishes.[4] This differential rate of change might reflect a real difference between the genders in response to the public health message. However, a recent analysis of the proportion of ever smokers who have quit (quit ratio) in the United States[4] suggests that the explanation lies in differences in age-specific cigarette smoking levels that result from the historical patterns of smoking uptake, i.e., the cohort effect suggested by Warner and Murt.[17]

For both males and females in the United States, the decline in smoking prevalence between 1974 and 1985 oc-

TABLE 6—International Trends in Smoking Prevalence in Highest and Lowest Educational Categories, 20 Years of Age and Over

	United States		Canada		Australia		Norway	
Year	Highest	Lowest	Highest	Lowest	Highest	Lowest	Highest	Lowest
1973								
1974	28.3	36.5					(29.4	45.8
1975			31.2	37.6				
1976	27.4	35.8			30.0	37.0		
1977	25.6	35.8	30.2	36.1				
1978	23.8	35.3						
1979	23.4	34.9	27.5	34.3				
1980	24.6	35.5			29.3	37.2		
1981			24.3	33.1				
1982								
1983	19.9	34.7	22.7	31.1	27.5	35.2		
1984								
1985	18.4	34.1					(22	43.5
1986			19.9	31.6	24.3	32.0		
1987	16.3*	35.7*						
Change year	−0.91**	−0.10**	−1.10‡	−0.61‡	−0.57‡	−0.50‡	−0.57‡‡	−0.18‡‡

*provisional data only
**obtained from weighted regression analysis of age-standardized data from 1974 to 1985[6]
‡obtained from simple linear regression of yearly prevalence data
‡‡obtained from averaging the change between the midpoints of each time intervals over the years
SOURCES: United States: Pierce, et al.[6]
Australia: Hill, personal communication
Canada: Millar[20]
Norway: Lund, personal communication

curred at a constant rate and the linear model closely predicted the 1987 prevalence. The decline in prevalence for each sex could very well be linear also in Great Britain, Canada, Norway, and Sweden. While caution is needed in comparing smoking prevalences in any given year, the United States is at the low end of any comparison. The country with by far the lowest prevalence is Sweden, although this might be confounded by the substitution of other tobacco products such as moist snuff, cigars, and pipes.* Of all the sex-specific trends reported, only women in Norway demonstrated an increase in smoking prevalence.

The rate of change in prevalence was much higher in Great Britain than in other countries, although this was also associated with a high initial prevalence in that country. Although methodological differences prevent precise statistical comparison, it is probable that the rate of change was higher for both sexes in Canada than in the United States and higher for women in Sweden. For both genders in Australia, although the linear model suggested that smoking prevalence had not declined much, the data are consistent with the reported large drop in prevalence that occurred with the introduction of the mass media-led coalition campaigns in that country.[18,19]

There also has been a consistent decline in the uptake of smoking among males in the United States in recent years. The rate of this decline was surpassed only in Canada and Sweden, although the estimates for the latter are somewhat unstable from small sample size. Again, in Australia, data indicated no linear trend, although this maybe an artefact of small sample size and an abnormally high reading for 1980. Uptake in young females in the United States did not decline between 1974 and 1985, but the 1987 prevalence suggests that such a decline has started to occur. This decline was well under way in other countries with large decreases in prevalence noted in Canada, Sweden, Norway, and Great Britain, although in each case prevalence started from a much higher base than in the United States. For both males and females, with the exception of Sweden, the United States had the lowest rate of uptake.

In every country smoking is much more prevalent in the lower socioeconomic groups than in the higher ones. Between 1974 and 1987, in both the United States and Canada, there was a major and consistent yearly decrease in smoking prevalence in the highest educational category. The rate of change was much greater in the highest compared to the lowest educational category in every country except Australia. Change in smoking prevalence did not start in Australia until 1983, after which it declined equally in both the highest and lowest categories.

In summary, there has been a major decline in smoking prevalence in all the countries studied. Comparison between the countries suggests that in recent times, the United States has had neither the lowest prevalence of smoking nor the fastest rate of decline in smoking. In particular, the rate of decline has been higher in Great Britain and Canada. Lack of change in prevalence among the lower educated is a major barrier to further marked decreases in smoking prevalence in the United States. This suggestion that public health action is not affecting a very sizable proportion of the population is apparent in all countries except Australia. In that country, with its State-based mass media-led campaigns, an equivalent decline was seen in both the highest and lowest educational categories since 1983. However, the lack of change in smoking prevalence in Australia throughout the decade before 1983 may reflect the lack of a clear government focus for smoking and its related health problems. Before 1983, any public health action on smoking came from the voluntary health agencies. The marked change since 1983 reflected the decision by a number of key state governments to act unilaterally and set up well-funded mass media-led antismoking campaigns. Australia still lacks a national bureaucratic structure whose sole responsibility is smoking and health. It

*Ramstrom L. Personal communication.

is apparent that all countries can benefit from further cross-fertilization of ideas and public health actions between countries. Undoubtedly, there is already considerable contact between public health professionals, from different countries concerned with reducing smoking in their own societies. However, the performance of each country in reducing smoking prevalence over the last 10 years indicates that each country has room for improvement in its public health action and suggests that interventions may not be diffusing across countries as rapidly as is either possible or desirable.

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APPENDIX
Educational Levels of Smokers in Five Developed Countries

Educational Level	United States	Austalia	Canada	Norway	Sweden
High					
Group 1	College graduate	University graduate	College graduate	University graduate	Highest third
Group 2	Some college	Post-secondary	Post-secondary		
Medium	High school graduate	Any secondary certificate	High School No post-secondary	Junior and Senior High	Middle third
Low	Less than high school graduate	Primary	Less than high school graduate	Primary	Lowest third

SOURCES: United States: Pierce, *et al*[5]
Australia: Hill, personal communication
Canada: Millar[12]
Norway: Lund, personal communication
Sweden: Ramstrom, personal communication

The Influence of Smoking on Vitamin C Status in Adults

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Abstract: To further define the relation between smoking and vitamin C status, the dietary and serum vitamin C levels of 11,592 respondents in the second National Health and Nutrition Examination Survey (NHANES II) were analyzed. Smokers of 20 cigarettes daily had the lowest vitamin C dietary intake (79 mg, 95% CI: 73, 84) and serum levels (0.82 mg dl, 95% CI: 0.77, 0.86; 46.6 μmol L, 95% CI: 43.7, 48.8), while smokers of 1–19 cigarettes daily had decreased vitamin C intake (97 mg; 95% CI: 90, 104 mg) and serum levels (0.97 mg dl, 95% CI: 0.92, 1.03; 55.1 μmol L, 95% CI: 52.2, 58.5) compared to respondents who had never smoked (109 mg, 95% CI: 105, 113 and 1.15 mg dl, 95% CI: 1.11, 1.18; 65.3 μmol L, 95% CI: 63.0, 67.0, respectively). This inverse association between both intake and serum levels of vitamin C and smoking was independent of age, sex, body weight, race, and alcoholic beverage consumption. Following further adjustment for dietary vitamin C intake, the negative correlation between cigarette smoking and serum vitamin C levels persisted. The risk of severe hypovitaminosis C (serum levels ≤ 0.2 mg dl; 11.4 μmol L) was increased in smokers, particularly when not accompanied by vitamin supplementation (odds ratio 3.0, 95% CI: 2.5, 3.6). These data suggest that even though smoking adversely affects preferences for vitamin C rich foods, the inverse association between smoking and serum vitamin C levels occurs independently of dietary intake. (*Am J Public Health* 1989; 79:158–162.)

Introduction

Preliminary reports have indicated adverse affects of smoking on vitamin C metabolism[1,2] and more recent studies have confirmed that serum vitamin C levels are lower in cigarette smokers than non-smokers.[3–7] These observations suggest that smoking directly lowers serum vitamin C levels,[8] although other factors including gender,[4,9] age,[4,10] and ethanol consumption[11–13] have also been reported to influence serum vitamin C concentrations. However, since cigarette smoking is not independent of age, sex, or alcohol use,[14] the relation between smoking and vitamin C status requires evaluation controlling for these potentially confounding variables. Several previous reports have partially addressed these concerns by stratifying participants according to age,[7,15,16] sex,[15,16] and level of vitamin C intake.[15] Within each stratum, vitamin C levels were decreased in smokers, suggesting that the relationship may indeed be independent of these factors. A univariate analysis of the United States second National Health and Nutrition Examination Survey (NHANES II) was consistent with these findings, as serum vitamin C levels were lower in smokers for each level of dietary vitamin C intake.[17]

The purpose of this study was to better define the independent relationship between cigarette smoking and vitamin C status, while controlling for potentially confounding variables using multivariate statistical techniques on data from NHANES II, conducted from 1976–80. NHANES II is a comprehensive population survey which included serum vitamin C measurements and a detailed dietary questionnaire from which nutrient intakes were calculated in a sample of over 11,000 adults. The association of cigarette smoking with serum vitamin C levels was analyzed while simultaneously adjusting for variables which significantly influenced serum vitamin C concentrations including age, sex, race, dietary vitamin C intake, ethanol consumption, and body mass index (BMI). In addition, the impact of smoking on the risk of marginal or frankly deficient serum vitamin C levels was determined.

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Methods

The NHANES II probability sample was selected so that population subgroups of special interest for nutritional assessment were oversampled. Among adults, the oversampled groups included persons over 60 years old and persons below poverty level. Sample weights were then computed for these subgroups so that the total sample closely estimated the noninstitutionalized civilian population. This sample weighting reflects the individual selection probability, adjustments for non-response, and poststratification adjustments. A complete description of the survey methods has been reported.[18]

Adults ages 18 to 74 years who attended a second examination visit following a preliminary interview were studied. However, respondents who did not have serum vitamin C concentrations measured (n=866), who did not record alcoholic beverage consumption (n=40), and who did not complete smoking histories (n=6) were excluded from the analysis. After these exclusions, 11,592 respondents remained eligible for the study. Appropriate sample weighting was used to adjust for possible non-response bias introduced because of failure to attend the second examination visit.[19] Those who attended the second visit but who were subsequently excluded (n=912) were similar to included subjects in the per cent who were male (46.6 vs 47.4 per cent), White (82 vs 89 per cent), currently smoking (32 vs 34 per cent), and in their mean age (46.7 vs 46.2 years).

Nutrient intakes were estimated from 24-hour recall interviews. Three-month food frequency reports provided recent intake of individual food items, including estimates for beer, wine, and liquor. Weekly frequencies of beer, wine, and liquor consumption were summed to estimate ethanol consumption. The use of vitamin and mineral supplements was determined by the question "Are you taking vitamins or supplements?" The vitamin content of supplements was excluded from the nutrient intake summaries and therefore the calculated vitamin C intake underestimates actual vitamin C consumption in those respondents taking supplements. For this reason, the use of supplements was included as a separate variable in the multivariate analysis assessing serum vitamin C levels.

NHANES II participants were originally divided into racial categories of "White", "Black", and "other". Individuals of Oriental extraction, American Indians, and Eskimos were listed in the "other" category and accounted for less than 1 per cent of the total study cohort; they were

included with Whites in this study. Body mass index (BMI) was measured as weight/height^2 (kg/m^2).

Serum vitamin C levels were measured in a central laboratory at the Centers for Disease Control (CDC) by the dinitrophenylhydrazine method.[20] Criteria from the Nutrition Canada Survey were used to classify risk of clinical scurvy.[21] Respondents with serum vitamin C levels of 0.2 mg/dl (11.4 μmol/L) or less were considered to be at high risk for clinical vitamin C deficiency (severe hypovitaminosis C), while those with serum concentrations of 0.4 mg/dl (22.7 μmol/L) or less were considered to be at marginal risk (marginal hypovitaminosis C).

Variance estimates were corrected for the complex sample design using the SUPERCARP program.[22] Multiple regression was used to determine the independent contribution of a variable to serum vitamin C levels, while simultaneously adjusting for all other variables.[23] An order of entry was determined using a stepwise regression model, however, regression coefficients were computed following the insertion of all variables into the model. For the regression equations, smoking status was categorized into six levels: never smoked, stopped for the previous year or longer (DC > 1 year), stopped within the past year (DC < 1 year), smoking less than 20 cigarettes daily (<1 PPD; mean 9.3 cigarettes), smoking 20 cigarettes daily (1 PPD), and smoking more than 20 cigarettes daily (>1 PPD; mean 35.1 cigarettes). This was done to compare serum vitamin C levels in cohorts of subjects who were currently smoking, had stopped smoking, or who had never smoked. In additional analyses, smoking status was also recorded as a continuous variable (i.e., number of cigarettes smoked daily). Odds ratios for smoking and nutrient supplements on the risk of hypovitaminosis C while controlling for possible confounding variables were determined by the Mantel-Haenzel procedure.[24,25]

Results

Table 1 shows univariate analyses for variables associated with both serum vitamin C levels and cigarette smoking prevalence. Serum vitamin C levels were significantly associated with gender, race, age, and body mass index. Males, when compared to females, Blacks when compared to Whites, and adults under the age of 60 years when compared with those over 60 had lower serum vitamin C levels and a higher prevalence of smoking. Heavier individuals (BMI > 27), on the other hand, had lower serum vitamin C concentrations and a lower prevalence of smoking. Higher ethanol intakes (>5 drinks daily) were associated with a higher likelihood of smoking, while a non-significant increase in vitamin C levels (p=0.09) was observed in those with more moderate consumption (1 to 5 drinks daily).

An inverse relation between cigarette smoking and dietary vitamin C intake is demonstrated in Table 2. The mean vitamin C intake for all smokers was 20% lower compared with non-smokers (86 mg, 95% CI: 83, 90 vs 108 mg, 95% CI: 106, 111 [not shown]). This decrease was particularly marked in individuals smoking 20 cigarettes daily. Smokers of less than 20 cigarettes daily had an intermediate vitamin C intake while individuals who stopped smoking for more than one year had a vitamin C intake (108 mg, 95% CI: 103, 113) similar to never smokers. This inverse association between smoking and dietary vitamin C intake persisted following adjustment for age, sex, race, ethanol intake, income and body weight.

TABLE 1—Dietary Vitamin C Intake, Serum Vitamin C Levels, and Smoking Frequency in Individuals by Category of Gender, Race, Age Tertiles, Body Mass Tertiles, and Alcohol Consumption

	N	Serum Vitamin C mg/dl* (95% CI)	Smokers (per cent)
Gender			
Male	5520	0.94 (0.90, 0.97)	40.5
Female	6072	1.12 (1.09, 1.16)	33.4
Race			
White	10430	1.05 (1.03, 1.08)	36.3
Black	1162	0.86 (0.80, 0.91)	41.2
Age (years)			
18–35	5178	1.14 (1.11, 1.18)	40.4
36–59	4614	1.01 (0.98, 1.04)	38.3
60–74	1800	1.02 (0.98, 1.06)	22.6
Body mass (BMI)			
Thin (<23)	4255	1.08 (1.05, 1.12)	43.1
Medium (23–27)	4090	1.05 (1.02, 1.08)	34.9
Heavy (>27)	3245	0.95 (0.91, 0.98)	30.9
Alcohol (drinks weekly)			
None	3875	1.02 (0.98, 1.05)	25.5
1–5	4509	1.06 (1.02, 1.10)	39.5
>5	1947	1.01 (0.96, 1.06)	50.1

*To convert to μmol/L, multiply by 56.78

In addition to decreased vitamin C consumption, smokers had serum vitamin C levels which were 24 per cent lower than non-smokers (0.87 mg/dl, 95% CI: 0.86, 0.89 vs 1.13 mg/dl, 95% CI: 1.12, 1.14; 49.4 μmol/L, 95% CI: 48.8, 50.5 vs 64.2 μmol/L, 95% CI: 63.6, 64.7 [not shown]). As shown in Table 3, smokers of 20 cigarettes daily had the lowest serum vitamin C concentrations (0.82 mg/dl, 46.6 μmol/L), while serum vitamin C levels were highest in individuals who had either stopped smoking for longer than one year or who had never smoked. The decrease in serum vitamin C levels observed in smokers persisted, even after adjusting for differences in dietary intake, age, gender, body weight, race, and vitamin supplementation. In multiple regression analysis, smoking status was the most important non-nutritive predictor of serum vitamin C levels, explaining more of the total variance than did gender, race, BMI, or age (Table 4). In additional analyses, smoking status was entered into the regression equation as a continuous variable defined as the number of cigarettes smoked daily. The order of entry and regression coefficients in these analyses were similar to those shown in Table 4 (i.e., when smoking was entered as a non-continuous variable).

The regression curves plotting dietary vitamin C intake

TABLE 2—Effect of Smoking Status on Dietary Vitamin C Intake

	N	Dietary Vitamin Intake (mg)	
		Mean (95% CI)	Adjusted Mean*
Smoker			
>1 PPD	1288	83 (76, 89)	79
1 PPD	1396	79 (73, 84)	78
<1 PPD	1498	97 (90, 104)	97
Non-smoker	—		
DC < 1 yr	335	97 (79, 116)	92
DC > 1 yr	1870	108 (103, 113)	105
Never smoked	4815	109 (105, 113)	111

NOTES: >1 PPD—Greater than 1 pack per day; 1 PPD—One pack per day; <1 PPD—Less than one pack per day; DC < 1 yr—Stopped smoking within the past year; DC > 1 yr—Stopped for the previous year or longer.

*Mean vitamin C intake adjusted for age, sex, race, ethanol intake, monetary income, and body weight.

TABLE 3—Effect of Smoking Status on Serum Vitamin C Levels

	N	Serum Vitamin C Concentration (mg/dl)†	
		Mean (95% CI)	Adjusted Mean*
Smoker			
1 PPD	1288	0.83 (0.79, 0.88)	0.89
<1 PPD	1454	0.82 (0.77, 0.86)	0.88
>1 PPD	1553	0.97 (0.92, 1.03)	0.99
Non-Smoker			
DC < 1 yr	346	0.99 (0.92, 1.06)	1.03
DC > 1 yr	1942	1.10 (1.06, 1.14)	1.11
Never Smoked	5009	1.15 (1.15, 1.18)	1.12

NOTES: >1 PPD—Greater than 1 pack per day. 1 PPD—Smoker of one pack per day. <1 PPD—Less than one pack per day. DC < 1 yr—Stopped smoking within the past year. DC > 1 yr—Stopped for the previous year or longer.

*Vitamin C serum level adjusted for age, sex, race, ethanol intake, body weight, dietary vitamin C intake, and supplementation status.

†To covert to μmol/L, multiply by 56.78.

with serum vitamin C levels for smokers and non-smokers are shown in Figure 1. The slope of the two lines is similar, with a consistent difference in serum vitamin C levels between smokers and non-smokers of approximately 0.2 mg/dl (11.4 μmol/L), independent of the level of dietary vitamin C intake. Using the regression equations, an additional 133 mg of dietary vitamin C was calculated to be necessary to increase the serum vitamin C level in smokers to the level found in non-smokers.

To test the hypothesis that smoking increases the risk of hypovitaminosis C, the frequency of low serum vitamin C levels occurring in smokers was compared with non-smokers. Of all smokers, 7.4 per cent had serum vitamin C levels of 0.2 mg/dl (11.4 μmol/L) or less compared with 1.9 per cent of non-smokers who had similar vitamin C levels. An additional 19.7 per cent of smokers had serum vitamin C levels between 0.2 mg/dl (11.4 μmol/L) and 0.4 mg/dl (22.7 μmol/L), compared with 8.2 per cent of non-smokers. Since further adjusting for BMI and age did not affect mean serum vitamin C levels, only gender, race, and dietary vitamin C intake were used as controlling variables to determine the risk of smoking on marginal or severe hypovitaminosis C. After adjusting for gender and race, risk ratios for smoking of 3.5 (95% CI: 2.9, 4.2) for severe, and 3.1 (95% CI: 2.8, 3.5) for marginal hypovitaminosis C were obtained (Table 5). Further adjustment for dietary vitamin C intake resulted in a minor reduction in the odds ratio.

Of the total sample, 37 per cent reported taking vitamin

TABLE 4—Stepwise Regression of Variables Influencing Serum Vitamin C Levels

Independent Variable	Order of Entry	Regression Coefficients	95% Confidence Intervals
Supplements	1	−0.3254	−.3005, −.3503
Smoking status	2	0.0515	.0467, .0564
Vit. C intake	3	0.0013	.0012, .0015
Sex	4	0.1379	.1200, .1558
BMI	5	−0.0098	−.0078, −.0118
Race	6	−0.1601	−.1181, −.2022
Age	7	0.0018	.0010, .0026
ETOH		0.0016	.0009, .0042

Total R^2: 0.29

Vitamin and mineral supplement use, coded as 1 = Yes, 2 = No.

Smoking status coded as non-continuous variable—see text.

Sex coded as 1 = male, 2 = female.

Race coded as 1 = White (and, other), 2 = Black.

ETOH is estimate of alcoholic drinks consumed weekly.

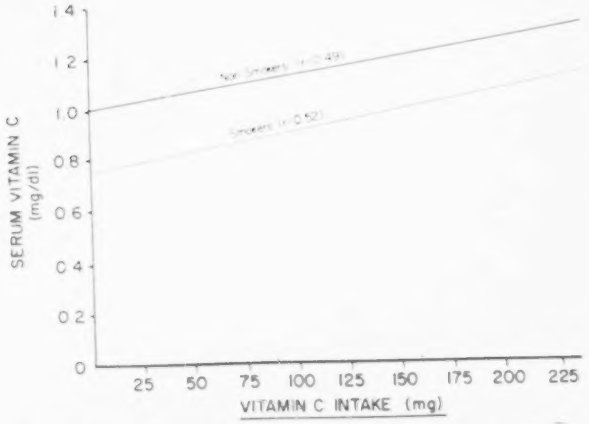

FIGURE 1—Linear Expression Curves for Smokers and Non-smokers Showing Relation between Dietary Vitamin C Intake and Serum Levels. (Adjusted for age, gender, race, body mass and vitamin supplementation. To convert serum levels to μmol/L, multiply by 56.78.)

or mineral supplements. The frequency of marginal hypovitaminosis C according to smoking and supplementation status is shown in Table 6. Smokers who did not take vitamin supplements had an 11.3 fold increased risk (95% CI: 9.6, 13.3) of marginal hypovitaminosis C compared with non-smokers who took supplements. Even smokers who took vitamin supplements were at increased risk of marginal hypovitaminosis C (odds ratio 2.5; 95% CI: 2.0, 3.1).

Discussion

Our analysis of the NHANES II data confirms that cigarette smoking is associated with decreased serum vitamin C levels. This association persisted despite correction for factors which independently affected serum vitamin C levels such as age, gender, race, and BMI. Dietary vitamin C intakes were lower in smokers than non-smokers, a finding which might potentially explain this inverse correlation. However, further adjustment for vitamin C consumption revealed that the association between smoking and serum levels was independent of dietary intake (Table 4). In fact, smoking status explained more of the total variance of serum vitamin C levels than did the self-reported dietary intake of vitamin C. Moreover, the increased risk of hypovitaminosis C in smokers was not substantially lowered when correcting for the decreased dietary vitamin C intake associated with smoking (Table 5). These observations suggest that cigarette smoking has an important influence on serum vitamin C levels which occurs predominantly via a mechanism independent of decreased dietary vitamin C consumption.

TABLE 5—Risk of Hypovitaminosis C in Smokers versus Non-smokers

	Unadjusted for Vitamin C Intake	Adjusted for Vitamin C Intake
Vitamin C Deficiency	Smoking Odds Ratio (95% CI)	Smoking Odds Ratio (95% CI)
Severe	3.5 (2.9, 4.2)	3.0 (2.5, 3.6)
Marginal	3.1 (2.8, 3.5)	2.8 (2.6, 3.1)

Unsupplemented subjects only.

TABLE 6—Combined Effect of Smoking and Vitamin Supplementation on Risk of Marginal Hypovitaminosis C

	Hypovitaminosis C Prevalence (%)	Risk of Hypovitaminosis C Odds Ratio (95% CI)*
Supplemented non-smokers	3.5	1.0
Supplemented smokers	9.8	2.5 (2.0, 3.1)
Unsupplemented non-smokers	14.5	3.9 (3.3, 4.7)
Unsupplemented smokers	35.7	11.3 (9.6, 13.3)

Inverse associations between smoking and vitamin C status have been previously reported. Most studies, however, have employed small sample sizes[1-5] or have been unable to control potentially confounding variables.[4,6,7] Pelletier stratified 4,500 Canadians participating in the Nutrition Canada National Survey into three categories of dietary vitamin C intake (0–30 mg, 30–100 mg, and over 100 mg) and analyzed serum vitamin C levels according to gender and age; an average 30 per cent reduction in serum levels were observed in smokers compared with non-smokers.[15] Ritzel, *et al.* determined that mean serum vitamin C levels were lower in smokers, independent of weekly citrus fruit consumption, in 4,053 healthy Swiss employees.[16] The risk of severe or marginal hypovitaminosis C was increased approximately threefold in smokers. Univariate analysis of NHANES II demonstrated a difference in serum vitamin C levels of 0.2 mg/dl (11.4 μmol/L) between smokers and non-smokers at each of six levels of dietary vitamin C intake.[17] These results are similar to the findings in this study using stepwise multiple regression to analyze the NHANES II data.

The lowered serum vitamin C levels in smokers could be due to either impaired vitamin C absorption or increased turnover. On the basis of studies measuring urinary vitamin C excretion in conjunction with the administration of known vitamin C intakes, Pelletier suggested the presence of impaired bioavailability but normal turnover of vitamin C in smokers.[26,27] On the other hand, Kallner, *et al.* measured vitamin C kinetics using radio-labeled ascorbic acid and demonstrated increased turnover in smokers but only small differences in absorption when compared to non-smokers.[28] Others have reported that smoking acutely increases urinary excretion of vitamin C,[29] also suggesting an accelerated metabolism in smokers. Interactions between vitamin C metabolism and a number of drugs have been described.[30,31] Identification of the responsible compound in tobacco smoke and the elaboration of its mechanism will require further study.

In this study, serum vitamin C levels were approximately 0.2 mg/dl (11.4 μmol/L) lower in smokers compared with non-smokers regardless of dietary vitamin C intake (Figure 1). Therefore, the impact of smoking to decrease serum vitamin C levels was proportionately greater in participants with initially low serum vitamin C levels. This resulted in a higher proportion of smokers with serum levels below the 0.2 mg/dl (11.4 μmol/L) or 0.4 mg/dl (22.7 μmol/L) range cutoffs for severe or marginal hypovitaminosis C. In this regard, the risk ratio for hypovitaminosis C was threefold greater in smokers when compared with non-smokers. Therefore, although the prevalence of severe and marginal hypovitaminosis C in NHANES II was low, certain subgroups, such as smokers not taking nutritional supplements, may be at substantial risk.

The significance of the increased risk of hypovitaminosis C produced by the interaction between smoking and serum vitamin C levels is not currently known. Although guidelines developed by the National Survey of Canada in 1973 suggested categories of high risk for serum vitamin C levels below 0.2 mg/dl (11.4 μmol/L), and of moderate risk for levels between 0.2 and 0.4 mg/dl [11.4 and 22.7 μmol/L],[21] the actual relation between low serum levels and clinical scurvy is not well defined. While serum levels in reported cases of scurvy are 0.4 mg/dl (22.7 μmol/L) or less,[32-34] individuals without clinical signs of scurvy may also have low serum levels.[35,36] However, because deficient vitamin C status has been associated with behavioral changes,[37,38] poor recovery from surgical procedures,[39] the presence of sublingual petechia,[40] and gingival disease,[41] and because the risk of overt scurvy is likely to be increased in this setting, the identification of individuals with hypovitaminosis C may have important implications. An assessment of the extent that hypovitaminosis C constitutes a true public health problem will require further studies to confirm that individuals with serum vitamin C levels below the cut-points used are truly at increased risk for adverse health outcomes.

The observed decrease in dietary vitamin C intake among smokers was unexpected. Although altered food preferences in smokers compared with non-smokers has been established,[42,43] a specific association of smoking with vitamin C intake is not well recognized. In the Caerphilly Heart Disease Study, smokers had a 22 per cent lower mean vitamin C intake compared with non-smokers,[44] similar to our findings in NHANES II. Our observation that vitamin C intake was lowest in individuals smoking 20 cigarettes or more daily, and that smokers who have stopped for greater than one year have intakes similar to individuals who have never smoked, suggest a specific aversive effect of smoking on vitamin C rich foods which may be reversible when smoking is discontinued. A more complete analysis from NHANES II of the effect of smoking on food preferences and vitamin intake will be reported separately.

Although an association between depressed vitamin C levels and smoking was determined to be independent of other potentially confounding variables, a cause and effect relationship cannot be assumed from this study. Prospective studies will be required both to prove that smoking directly affects vitamin C metabolism and to evaluate possible mechanisms. However, the cohort of participants who smoked less than 20 cigarettes daily had lower mean serum vitamin C levels than heavier smokers but higher than non-smokers, suggesting a dose-response effect. In addition, former smokers had higher serum levels than current smokers, indicating that the influence of smoking on serum vitamin C levels may be reversible. These observations suggest that the association between smoking and serum vitamin C levels may be causal.

This study is also limited by the methodological problems associated with estimating dietary intake in cross-sectional studies of free-living populations. Dietary recall of the preceding 24-hour period may not provide an accurate or representative assessment of current nutritional intake for an individual. However, this limitation is lessened when assessing the average intake of a group, since the group mean intake is less sensitive to day-to-day variability.[45] An additional shortcoming is that specific nutritional supplements were not systematically recorded and therefore were not included in the dietary analysis. This problem has been addressed in this study by either adjusting or stratifying for supplement use.

This analysis of NHANES II data confirms that cigarette smokers in the United States have lower serum concentrations of vitamin C than non-smokers, an association which persisted following adjustment for potentially confounding

factors. Because of these decreased levels, smokers had a three-fold increase in risk of marginally or severely deficient vitamin C status. Approximately 35 per cent of smokers had low serum levels placing them within one of these two categories. Although smokers taking vitamin supplements were also at increased risk of hypovitaminosis C compared with non-smokers, the prevalence of low serum vitamin C levels was much reduced in both smoking and non-smoking individuals taking supplements. This finding suggests that the increased vitamin C intake associated with vitamin supplementation may frequently prevent the low serum levels from occurring in smokers. As estimated by the regression curves in Figure 1, approximately 130 mg of additional dietary vitamin C daily would be required to overcome the adverse effect of cigarette smoking on serum vitamin C levels. Although vitamin C supplementation has been associated with altered protein and amino acid metabolism,[46] increased urinary oxalate excretion leading to calcium oxalate calculi,[47] diminished serum levels of vitamin B12,[48] and interference with urine tests for glucose and stool tests for occult blood,[49,50] these adverse effects have been reported only with daily vitamin C doses of one gram or more. Therefore, increasing the dietary vitamin C intake by an amount sufficient to correct the hypovitaminosis C associated with smoking should be well tolerated. Although the simplest and most direct method to increase the low serum vitamin C levels found in many smokers would be to stop smoking, increasing vitamin C consumption may be appropriate when cigarette cessation is unsuccessful.

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Deaths from All Causes in Non-Smokers Who Lived with Smokers

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Abstract: Mortality associated with passive smoking was evaluated in a 12-year study of 27,891 White adult smokers and 19,035 never smokers identified in 1963. Death rates were calculated using an estimate of the person-years at risk. Adjusted for age, marital status, education, and quality of housing, the estimated relative risks of death from all causes were 1.17 (approximate 95% confidence interval 1.01, 1.36) for men and 1.15 (1.06, 1.24) for women with passive exposure. These relative risks were similar to those for ex-smokers and for pipe or cigar smokers. Risks increased slightly with level of exposure. The relative risk from passive smoking was greatest for men under age 50 (RR = 2.09, 1.31–3.34). Risks from passive smoking were slightly elevated for several causes among men and women, and may be broader than those previously reported. On the other hand, these small nonspecific increases in death rates may reflect other characteristics of passive smokers that increase mortality. (*Am J Public Health* 1989; 79:163–167.)

Introduction

Smokers are generally reported to have almost doubled age-adjusted death rates.[1] This increased mortality is observed for all causes combined as well as for specific causes, and is not entirely explained by smokers' substantially increased risk for lung cancer or arteriosclerotic heart disease.

Non-smokers who live with smokers are exposed to many of the same potentially toxic consituents of cigarette smoke[2] and they, too, may be at increased risk for death from a variety of causes. These "passive smokers," like active smokers, have increased levels of smoke by-products in body fluids,[3,4] increased activation of enzyme systems that metabolize potential carcinogens,[5] and increased levels of urinary mutagens,[6,7] albeit all at much lower levels than active smokers.

More than a dozen studies have demonstrated increased lung cancer risk among passive smokers,[2] and several studies have suggested increased heart disease risk for passive smokers.[8–11] Studies have also reported increased overall cancer risk,[12–14] with risks not limited to respiratory tract cancers, but the data are limited and findings inconclusive. Only a few reports have concerned total mortality among passive smokers. In the Multiple Risk Factor Intervention Trial, total mortality was increased nearly two-fold among men at risk for heart disease who were married to smokers.[8] A small increase in total mortality was reported for non-smoking women with environmental tobacco smoke exposure in Scotland, but no association was observed for men.[15] Total mortality was slightly increased for passively exposed women in a small US study,[9] but no association between passive smoke exposure and total mortality was found in a study in Amsterdam.[16]

We evaluated mortality from all causes and specific causes in non-smokers living with smokers who had been surveyed in 1963 as part of a private community-wide census in Western Maryland.

Methods

In July 1963 a private census obtained data on an estimated 98 per cent of the households in Washington County, Maryland. The census provided information on sex, age, race, marital status, years of schooling, church attendance, and housing characteristics for 91,909 individuals. Information on cigarette, cigar, and pipe smoking habits was recorded for each household member age 16-½ or older as of July 15, 1963.

To take into account losses from the population resulting from deaths and emigration, a follow-up study of a 5 per cent sample of households that took part in the 1963 census was conducted in 1971. Age, marital status, years of schooling, and frequency of church attendance were associated with remaining in the county after eight years based on a linear model that adjusted for effects of these and other socioeconomic and demographic variables. The addition of sex, smoking status in 1963, and number of bathrooms to the model did not alter the results. The resulting linear model was used to assign to each adult in the 1963 census their probability of still living in the county. From this, an annual probability could be calculated, allowing the population remaining in the county to be estimated at any point in time. The procedure used to estimate the probability of remaining in the county has been described in greater detail by Comstock and Tonascia.[17]

Death certificates of Washington County residents who died between July 1963 and July 1975 were matched against the census and have been coded by one of the authors (GWC) as to primary, contributing, and underlying causes of death according to the seventh revision of the International Classification of Disease (ICD).[18] For this report, deaths were classified in broad groupings using the underlying cause of death.

Data are reported here for adults who were over age 25 in 1963. Because only about 2 per cent of the census population was Black, the analysis is limited to Whites. This report is based on data from a total of 22,973 White men and 25,369 White women.

Study subjects were classified according to their smoking status in 1963. Smoking status was not known for 587 men (2.6 per cent) and 527 women (2.1 per cent). An additional 212 men and 90 women known to be smokers were excluded from this report because the number of cigarettes smoked was not known, leaving 22,174 men and 24,752 women for analysis.

A household smoke exposure score was calculated to serve as a measure of passive smoking. To create this score, each adult was assigned a smoking contribution score ranging from 0 to 12 based on their personal smoking history.[11] A household smoking total was calculated as the sum of the smoking contribution scores of all persons living in that household, and each individual's household smoke exposure

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TABLE 1—All Cause Mortality by Smoking Status in 1963

Smoking Status	Men				Women			
	N#	Deaths	Adj RR*	(95% CI)	N#	Deaths	Adj RR*	(95% CI)
Non-smokers	2,914	585	1.00	—	5,322	1,224	1.00	—
Passive smokers	1,248	258	1.17	(1.01, 1.36)	9,551	1,556	1.15	(1.06, 1.24)
Pipe cigar only	1,671	504	1.20	(1.07, 1.35)	10	2	0.70	(0.18, 2.75)
Ex-smokers	4,562	997	1.16	(1.05, 1.28)	2,153	186	1.03	(0.87, 1.21)
Current smokers	11,779	2,090	1.41	(1.29, 1.55)	7,716	697	1.26	(1.13, 1.40)
<10 cigarettes	1,583	315	1.14	(0.99, 1.31)	2,248	222	1.12	(0.96, 1.30)
10+ cigarettes	10,196	1,775	1.49	(1.35, 1.64)	5,468	475	1.34	(1.19, 1.52)

Number of subjects
*Adjusted for age, housing quality, schooling, and marital status

score was the household total minus his or her own contribution to it. The person's own smoking was treated separately in the analysis.

The household exposure score cannot measure total passive smoke exposure because it is not known, for example, how many of the cigarettes smoked per day were smoked outside of the home. Nevertheless, this score is an indicator of the relative degree of exposure in 1963. The group of non-smokers living with non-smokers will have had, on average, the least household exposure to tobacco smoke, while persons living with several heavy smokers will have had, on average, the highest levels of household exposure. Among the 4,162 men and 14,873 women who were reported in 1963 as never smokers, 1,248 men (30.0 per cent) and 9,551 women (64.2 per cent) had household smoke exposure and are considered passive smokers; the remainder are considered non-smokers.

A housing index (ranging from 0 to 10) based on running water, number of bathrooms, type of heating system, cooking fuel, and availability of a telephone was created to serve as a rough indicator of quality of housing and socioeconomic status. For this analysis, housing quality has been classified as poor (index = 0–7) or good (index = 8–10). Education has been categorized as 0–8, 9–11, 12, and 13 or more years of schooling. Subjects were characterized as those married in 1963 and those not. In overall analyses, age has been categorized as 25–44, 45–54, 55–64, and 65+. In analyses restricted to those under age 50, age was categorized as less than 40 and 40–49 in 1963. For persons older than 50, age was categorized as 50–59, 60–69, and 70 or more years.

Average annual death rates were calculated as the number of deaths during the 12 years from 1963 to 1975 divided by the number of person-years of follow-up estimated from the assigned probability scores for still living in the county. Relative risks were calculated as the average annual death rates among smokers or passive smokers divided by the death rates among nonsmokers without household smoke exposure. Age, marital status, and other factors were strongly related to both death rates and smoking status in this population. Accordingly, all relative risks were adjusted for differences in age, marital status, housing quality, and education using Poisson regression.[19]

Results

The number of deaths and person years of follow-up are shown by smoking status in 1963 and sex in Table 1. After adjusting for age, marital status, education, and housing quality, passive smokers were at increased risk for death from all causes relative to non-smokers. Adjusted relative risks were 1.17 (approximate 95 per cent confidence intervals 1.01, 1.36) for men and 1.15 (1.06, 1.24) for women. These relative risks were similar to those for pipe or cigar smokers, ex-smokers, and smokers of fewer than 10 cigarettes a day, and were less than half that observed for heavier active smokers.

Among men, the relative risk for passive smokers under age 50 was 2.09 (1.31, 3.34), and the relative risk for heavier active smokers was 3.16 (2.27, 4.41) (Table 2). Among those age 50 and older, the corresponding relative risks were substantially less. The difference between the death rates for passive smokers and non-smokers was similar, however, in both age groups (2.6/1,000 for those younger than age 50 and 2.3 for those 50 and older). Among women, while the relative risk for active smoking was greater for those younger than 50 in 1963, the relative risk for passive smoking did not vary with age. Among passive smokers there was some increase in

TABLE 2—All Cause Mortality by Smoking Status and Age in 1963

Smoking Status	Men				Women			
	Age <50		Age 50+		Age <50		Age 50+	
	Adj RR*	(95% CI)	Adj RR	(95% CI)	Adj RR	(95% CI)	Adj RR	(95% CI)
Non-smokers	1.00	—	1.00	—	1.00	—	1.00	—
Passive smokers	2.09	(1.31, 3.34)	1.11	(0.95, 1.30)	1.14	(0.81, 1.62)	1.14	(1.05, 1.23)
Pipe cigar only	1.83	(1.13, 2.96)	1.13	(1.00, 1.28)	—	—	0.72	(0.18, 2.86)
Ex-smokers	1.94	(1.34, 2.81)	1.15	(1.03, 1.28)	1.57	(1.05, 2.36)	1.00	(0.83, 1.20)
Current smokers	3.06	(2.20, 4.27)	1.33	(1.20, 1.47)	1.79	(1.30, 2.47)	1.27	(1.13, 1.43)
<10 cigarettes	2.28	(1.49, 3.48)	1.12	(0.97, 1.30)	1.34	(0.88, 2.02)	1.18	(1.00, 1.39)
10+ cigarettes	3.16	(2.27, 4.41)	1.39	(1.26, 1.55)	1.96	(1.42, 2.71)	1.33	(1.16, 1.53)

*Adjusted for age, housing quality, schooling, and marital status

TABLE 3—All Cause Mortality among Never Smokers by Degree of Passive Smoke Exposure and Age

Exposure Score	Age < 50			Age 50+			Total		
	Deaths	Adj RR*	(95% CI)	Deaths	Adj RR	(95% CI)	Deaths	Adj RR	(95% CI)
Men									
0 (none)	37	1.00	—	548	1.00	—	585	1.00	—
1+ (any)	35	2.09	(1.31, 3.34)	223	1.11	(0.95, 1.30)	258	1.17	(1.01, 1.36)
1–5 (light)	12	1.59	(0.83, 3.07)	95	1.11	(0.89, 1.38)	107	1.13	(0.92, 1.39)
6+ (heavy)	23	2.50	(1.48, 4.24)	128	1.11	(0.91, 1.35)	151	1.20	(1.00, 1.44)
Women									
0	47	1.00	—	1,177	1.00	—	1,224	1.00	—
1+	132	1.14	(0.81, 1.62)	1,424	1.14	(1.05, 1.23)	1,556	1.15	(1.06, 1.24)
1–5	53	1.24	(0.83, 1.86)	650	1.09	(0.98, 1.21)	703	1.12	(1.02, 1.24)
6+	79	1.08	(0.75, 1.58)	774	1.17	(1.06, 1.28)	853	1.16	(1.06, 1.27)

*Adjusted for age, housing quality, schooling, and marital status.

overall relative risk for death from all causes with increasing household exposure to smoke, but no clear trends (Table 3).

Passive smoking-associated adjusted relative risks for individual causes of death are shown for never smoking men and women in Table 4. The relative risks for male and female passive smokers were greater than one for several smoking related and other causes, although risk estimates were stable only for arteriosclerotic heart disease in men and for arteriosclerotic heart disease, cerebrovascular disease, and chronic lung disease in women.

There were few deaths among men younger than 50 in 1963, but the relative risks for several causes of death, including arteriosclerotic heart disease and cancer, were greater than 2.00 (data not shown).

The increased relative risk for chronic lung disease among women with smoke exposure was limited to women over age 50 in 1963; there were no deaths in this category among younger women. The risk for cerebrovascular disease did not differ by age, but cancer risk was increased only among women younger than age 50. All 12 "smoking-related" cancers (two lung, five cervix, one kidney, three pancreas) among never smoking women younger than 50 were in those with passive smoke exposure.

Among the never smokers, 44 per cent of the deaths among men and 35 per cent of the deaths among women were attributed to arteriosclerotic heart disease. After excluding these deaths, there continued to be an increase in overall mortality for women with passive smoke exposure (Table 5). For men, mortality from causes other than arteriosclerotic heart disease was increased only among passive smokers younger than age 50 [RR = 1.93 (1.09, 3.41)].

Discussion

The finding of increased overall mortality for non-smokers who lived with smokers may not reflect a causal relationship. Much of the increased risk we observed could be attributed to arteriosclerotic heart disease for which an extensive literature documents an increased risk associated with active cigarette smoking.[20] Our own study[11] and others have demonstrated increased risk for heart disease in passive smokers. Svendsen, *et al.*[8] reported non-smoking men married to smokers were at increased risk for heart disease death [RR = 2.11 (0.69, 6.46)] and risk increased with amount smoked by the spouse. Hirayama[10] reported a 20 per cent increase in mortality from ischemic heart disease among non-smoking women married to smokers. Heart disease

TABLE 4—Deaths from all Causes and Specific Causes among Never Smoking Men and Women

Cause of Death	Men				Women			
	Deaths				Deaths			
	Passive Smokers (12,238 PY)	Non-smokers (29,214 PY)	Adj RR*	(95% CI)	Passive Smokers (97,032 PY)	Non-smokers (51,113 PY)	Adj RR	(95% CI)
All causes	258	585	1.17	(1.01, 1.36)	1,556	1,224	1.15	(1.06, 1.24)
Cancer	31	84	1.01	(0.66, 1.53)	290	211	1.00	(0.82, 1.21)
Smoking related	8	24	0.96	(0.43, 2.16)	49	27	1.45	(0.88, 2.40)
Other	23	60	1.03	(0.40, 2.62)	241	184	0.93	(0.76, 1.54)
Cerebrovascular disease	33	93	0.97	(0.65, 1.46)	297	232	1.24	(1.03, 1.49)
Circulatory diseases	137	304	1.19	(0.97, 1.46)	715	576	1.17	(1.05, 1.32)
Arteriosclerotic heart disease	122	248	1.31	(1.05, 1.64)	551	437	1.19	(1.04, 1.36)
Other	15	56	0.65	(0.36, 1.16)	164	139	1.14	(0.89, 1.29)
Respiratory diseases	15	27	1.44	(0.75, 2.75)	79	72	1.21	(0.87, 1.69)
Emphysema Bronchitis	2	4	0.93	(0.16, 5.32)	11	2	5.65	(1.19, 26.8)
Other	13	23	1.53	(0.76, 3.07)	68	70	1.08	(0.76, 1.54)
Digestive diseases	9	18	1.34	(0.59, 3.04)	57	39	1.20	(0.78, 1.85)
Genitourinary diseases	3	12	0.67	(0.36, 1.28)	21	11	1.50	(0.69, 3.26)
Accidental causes	15	24	1.48	(0.78, 2.84)	48	35	1.24	(0.78, 1.30)
Other	15	23	1.70	(0.88, 3.29)	49	48	0.76	(0.50, 1.17)

*Adjusted for age, housing quality, schooling, and marital status.

TABLE 5—Mortality among Never Smokers from Arteriosclerotic Heart Disease and Other Causes

Cause of Death	Men		Women	
	Adj RR*	(95% CI)	Adj RR	(95% CI)
Arteriosclerotic heart disease	1.31	(1.05, 1.64)	1.19	(1.04, 1.36)
Other causes	1.06	(0.87, 1.31)	1.12	(1.02, 1.24)

*Adjusted for age, housing quality, schooling, and marital status.

mortality associated with passive smoking was also increased in studies reported by Gillis, *et al.*[15] and by Garland, *et al.*[9] but risk estimates were based on small numbers. A study reported by Lee, *et al.*[21] indicated no heart disease risk from passive smoking.

In our study, an increased relative risk for death remained even after subtracting deaths from arteriosclerotic heart disease. The effect was strongest for women and for younger men. The general increase in mortality leaves open the possibility that the life styles of people who live with smokers differ from those who do not live with smokers. Factors such as alcohol consumption or dietary habits which are correlated with both smoking and risk for some diseases (and which were not collected in the 1963 census) seem especially likely to be alternative explanations for our findings, to the extent that diets or alcohol use are similar among household members.

Studies of passive smokers in general and of passive smokers with heart disease have also examined some of these potential confounding factors with mixed results. In 1963, more than 60 per cent of never smoking women in Washington County lived with smokers,* but it was much less common for non-smoking men to live with smokers. Differences between these men and other men might explain their slight increase in overall mortality. Never smokers who lived with smokers had fewer years of schooling and lived in housing of poorer quality. We adjusted for these differences in our analysis, but residual confounding might still affect our results.

Others have found few differences between those with and without passive smoke exposure that could account for mortality differences between groups. Garland, *et al.*[9] found no difference in blood pressure, obesity, or cholesterol between women married to smokers and those married to never smokers. Svendsen, *et al.*[8] found no differences in blood pressure, cholesterol, or psychosocial factors between men married to smokers and men married to non-smokers. They did, however, find that men married to smokers were slightly heavier and drank alcoholic beverages slightly more often than those married to non-smokers. Similarly, Friedman, *et al.*[22] reported that alcohol consumption and other risk-taking behaviors were more frequent among non-smokers with passive smoke exposure.

Besides heart disease, several other causes of death for which we observed increased death rates for passive smokers are causes that are known or suspected to be caused in part by cigarette smoke. These include respiratory diseases such as emphysema and bronchitis[23] and cerebrovascular disease.[24,25]

Kalandidi, *et al.*[26] recently reported in a case-control study that non-smoking women married to current smokers had nearly twice the risk of lung disease as did non-smoking women married to never smokers. Increased chronic lung disease mortality was also reported by Hirayama[27] for Japanese women married to smokers (RR = 1.4), but results in the study by Lee, *et al.*[21] were inconclusive. Two studies also included data on risk of cerebrovascular disease. Based on very small numbers, stroke risk was increased for women with environmental tobacco smoke exposure in the study by Gillis, *et al.*[15] In the study by Lee, *et al.*[21] stroke mortality was increased approximately two-fold for both men and women with high total passive smoke exposure, although not in relation to spouse smoking alone.

While lifestyle or selection factors are a possible explanation for the small increase in overall mortality we observed, it is also possible that continued exposure to cigarette smoke leads to poorer health in general. Smokers have more respiratory illnesses and decreased lung function relative to non-smokers.[24] They may also have altered immune system responses and other changes which indirectly affect disease risk and response.[1] It is possible that passive smoke exposure indirectly increases overall mortality by altering general health status and increasing case-fatality from other illnesses. Increased respiratory symptoms and decreased lung function have been observed for passively exposed children and adults,[2] although not all studies have been positive.[28]

Finally, with relative risks as low as those observed here, and with such a weak dose-response, one must consider that other biases are responsible for these findings. The confidence intervals for the risk estimates shown do not take into account the variability associated with the estimated person years at risk. These estimates were derived from fitting a linear model to follow-up data obtained for a 5 per cent sample of the population. The true person-years of follow up may be larger or smaller than that calculated. Therefore the reported confidence intervals may be slightly too small.

Misclassification of smoking status is also possible. Smoking status was obtained once in 1963 as part of a general community census. Some persons classified as non-smokers may have taken up smoking in the years between 1963 and 1975, and non-smokers who lived with smokers may be more apt to become smokers. Similarly, non-smokers with no smoke exposure may have become exposed as other family members began to smoke or as household composition changed. However, it is more likely that active smokers became ex-smokers over the 12-year period, reducing the overall level of household smoke exposure at which these effects were observed. Thus the dose-specific effects may even be underestimated. Furthermore, individuals who were not exposed at home may have had smoke exposure at work, making it difficult to detect smoking-related health differences between groups.

*Sandler DP, Helsing KJ, Comstock GW, Shore DL: Factors associated with past household exposure to tobacco smoke. Submitted for publication.

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NIDA Announces Availability of Videotapes on Drugs in the Workplace

The National Institute on Drug Abuse, the federal agency which serves as the focal point for research and demonstration programs on drug abuse, has announced availability of *Drugs at Work* and *Getting Help*—the first two in a four-part videotape series on drug abuse in the workplace. The videotapes are available either on free loan or for purchase.

- *Drugs At Work*, a 23-minute educational documentary, describes costs of drug use for the workplace, the individual, and the public, and examines action being taken by government and private companies. An important program for all levels in the workplace, from top management to line staff, *Drugs at Work* provides information for developing effective drug abuse workplace programs through policy development, drug testing, employee assistance programs, prevention, and education. Interviews are included with drug users who have sought treatment and with experts on drugs in the workplace. Government and industry representatives describe current federal and corporate programs.
- *Getting Help* presents detailed information about the use of Employee Assistance Programs (EAPs) in addressing drug use in the workplace. The film features comments by business, labor, and government leaders, and EAP professionals; presentation of three model programs; and EAP client interviews. It encourages employers to consider EAPs as a tool in combatting drugs at work, and provides employees with reassuring information about the confidentiality and effectiveness of an EAP program.

The third in the series, *Drug Testing*, and the fourth, *Drug Prevention and Employee Education*, will be available in the near future.

The videotapes—available in both employer and employee versions—are designed to be part of a comprehensive workplace drug abuse program. To arrange a free loan, contact: Modern Talking Picture Service Scheduling Center, 5000 Park Street, North, St. Petersburg, FL 33709, (813) 541-5763. To purchase, contact: National Audio Visual Center, Customer Service Section, 8700 Edgeworth Drive, Capital Heights, MD 20743–3701, (301) 763–1896.

Parental Smoking and Post-Infancy Wheezing in Children: A Prospective Cohort Study

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Abstract: The contribution of parental smoking to wheezing in children was studied in a subset of all British births between April 5 and 11, 1970 ($N = 9,670$). Children of smoking mothers had an 18.0 per cent cumulative incidence of post-infancy wheezing through 10 years of age, compared with 16.2 per cent among children of nonsmoking mothers (risk ratio 1.11, 95% CI: 1.02, 1.21). This difference was confined to wheezing attributed to wheezy bronchitis, of which children of smokers had 7.4 per cent, and those of nonsmokers had 5.2 per cent (risk ratio 1.44, 95% CI: 1.24, 1.68). The incidence of wheezy bronchitis increased as mothers smoked more cigarettes. After multiple logistic regression analysis was used to control for paternal smoking, social status, sex, family allergy, crowding, breast-feeding, gas cooking and heating, and bedroom dampness, the association of maternal smoking with childhood wheezy bronchitis persisted. Some of this effect was explained by maternal respiratory symptoms and maternal depression, but not by neonatal problems, the child's allergic symptoms, or paternal respiratory symptoms. There was a 14 per cent increase in childhood wheezy bronchitis when mothers smoked over four cigarettes per day, and a 49 per cent increase when mothers smoked over 14 cigarettes daily. (*Am J Public Health* 1989; 79:168–171.)

Introduction

Children are passive victims of the effects of tobacco smoke from their surrounding environment, resulting in multiple adverse health outcomes.[1–3] Passive smoking has been associated with respiratory infections and symptoms in infants and young children.[4–9] Some studies have suggested an association between parental smoking and wheezing or other respiratory symptoms of older children,[10–14] but other investigations have disputed this relationship.[15–17] A dose effect has been reported,[4,6,12,18–21] although this has been inconsistent and often crudely measured.[8] Parental smoking has also been associated with decreased pulmonary function and lung growth in children.[11,14,19,22–25]

The mechanism of effect of passive smoking on childhood reactive airway disease is not clear. Sidestream tobacco smoke may be directly toxic to the lungs of children. Alternatively, or additionally, the respiratory effects of passive smoking may be mediated by other factors to which children of smokers may be exposed, such as infection,[10,13,26,27] allergy,[15,28] stress,[29] or neonatal respiratory problems.[23]

In order to address some of these unresolved issues, we studied the relationship of cigarette smoking by either or both parents with report of post-infancy wheezing in a large, representative British national cohort.

Methods

The Child Health and Education Study began as a national survey of all births in Great Britain during the week of April 5–11, 1970.[30] Information was collected by midwives during the first postpartum week. At 5 years of age, the cohort was reevaluated by trained health visitors via maternal interviews. Additional parental interviews, medical examinations, and educational evaluations were carried out at age 10.

Maternal smoking behavior was surveyed at each interview, including the duration of smoking from pregnancy through age 10 of the child and the usual dose of cigarettes per day during each period. We asked about the use of cigarettes, pipes, or cigars by the father. The amount smoked by the mother was categorized on the birth questionnaire as 0, 1–4, 5–14, 15–24, and 25 or more cigarettes per day. These categories were applied to doses reported at children's ages 5 and 10, from the continuous measures of amount smoked (there was considerable preference for multiples of five in reporting of dose). In trend analyses, means of each of the first three smoking categories were used: 2.5, 9.5, and 19.5; 29.5 cigarettes per day was used for the highest smoking category.

Outcome data were obtained from maternal interviews when the children were 10 years old. At that time, parents were asked: "Has the child *ever* had one or more attacks or bouts in which there was wheezing or whistling in the chest?" If the reply was positive, they were asked: "What were these thought to be due to?" (. . . asthma, wheezy bronchitis, or other cause, with multiple responses permitted); parents were then asked at what ages the wheezing occurred. Wheezy bronchitis is a common diagnosis in Great Britain, although its clinical distinction from asthma is not clear.

Other variables were selected for study after review of published investigations or parental smoking and other factors related to childhood respiratory symptoms and disease. These included sex of the child, social status of the family (represented by a standardized variable derived from factor analysis of social items when the cohort was 5 years old), crowding in the home (total persons/total rooms), number of siblings (total, older and younger), dampness in the child's bedroom, whether gas was used for cooking and/or heating, child's own smoking at age 10, whether the child was breast fed at least one month, history of allergy in the child or family, neonatal respiratory problems, symptoms of chronic cough or phlegm for over three months in each parent, and a measure of maternal depression (derived from a 24-item Malaise Inventory[31]). Family allergic history was determined at the time the child was age 5 by a report of asthma, eczema, or hay fever in either natural parent. A positive allergic history in the child was defined as a report of eczema or hay

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fever before 5 years of age. Neonatal respiratory problems were defined by a history of any breathing difficulties during the first week of life.

Chi-square, multiple logistic regression, and factor analyses were done using SAS software.[32,33] Confidence intervals on risk ratios[34] and on odds ratios determined from multiple logistic regression[35] were calculated using published methods. Attributable risk calculations were done using methods described by Lilienfeld.[36]

Results

We studied prospectively 11,776 children born between April 5–11, 1970 and their families, obtaining data at birth, and at ages 5 and 10 years. Of this population, 9,953 (84.5 per cent) were singleton births, living throughout with their natural mothers, whose primary language was English, and who were ethnic Europeans. In 836 of these households (8.4 per cent), there was no father figure. There were 5,113 boys (51.4 per cent).

The 9,670 children whose mothers had known smoking information comprise the cohort for the remaining analyses. Of these mothers, 4,855 (50.2 per cent) smoked cigarettes at some time between their pregnancy and the child's age 10. Cigarette smoking was reported at some time by 5,128 fathers in homes with a father figure (56.2 per cent). In addition, 866 (9.5 per cent) of fathers smoked pipes or cigars only.

There were 1,774 children (17.8 per cent) who were reported to have wheezed by 10 years of age. Of these, 112 reported wheezing only before one yer of age; the remaining 1,662 were classified by their parents as having asthma, wheezy bronchitis, or other reasons for wheezing. The relationship of post-infancy wheezing to maternal smoking at any time between pregnancy and child's age 10 (dichotomized) is presented in Table 1. Smoking was associated with post-infancy wheezing in the entire group of children (risk ratio 1.11, 95 per cent CI = 1.02, 1.21). This smoking relationship was limited to the subgroup stated to have had wheezy bronchitis (risk ratio 1.44, 95 per cent CI = 1.24, 1.68). Subsequent analyses relate only to this outcome.

Unadjusted relationships of maternal and paternal smoking with child's wheezy bronchitis are shown in Table 2. The effect of maternal smoking was stronger when fathers also smoked, but these mothers also smoked more cigarettes per day: during pregnancy, 12.5 per cent smoked over 14 cigarettes per day, compared with 4.5 per cent when fathers were nonsmokers. In the homes with no father figure, there was a higher cumulative incidence of wheezy bronchitis.

Wheezy bronchitis was reported to have started before age 5 in 72 per cent of affected children (compared with 65 per cent of those with asthma). The amounts that the mother smoked at child's birth and age 5 years were equally strongly related to the risk of wheezy bronchitis. Thus, the mean dose of cigarettes per day between the reported amounts at birth and 5 years was used in subsequent dose analyses. The cumulative incidence of wheezy bronchitis by amount smoked by the mother is presented in Table 3. The incidence of wheezy bronchitis increased progressively from 5.2 per cent in nonsmokers to 8.9 per cent among children whose mothers smoked over 24 cigarettes per day, but this dose-effect trend among smokers was not statistically significant.

Wheezy bronchitis was inconsistently related to the duration and period of maternal smoking (Table 4). Those mothers who smoked throughout the study period (from pregnancy through child's age 10) also smoked more cigarettes; their children had a risk ratio of 1.52 (95 per cent CI = 1.27, 1.82) for wheezy bronchitis compared with nonsmokers. The category of mothers who only smoked in both postnatal periods was most strongly related to childhood wheezy bronchitis, although only 12 cases were exposed in this group.

Multiple logistic regression was used to estimate the effects of several factors of post-infancy wheezy bronchitis contrasted with no history of wheezing. The final model included maternal smoking between child's birth and age 5 years with four dose levels; paternal smoking in three categories: cigarette smoking, pipe or cigar smoking only, or father absent from the home; social status, sex of the child, family history of allergy, crowding, breast-feeding, gas cooking or heating, and bedroom dampness. The included social status items after factor analysis were: father's highest educational qualification, father's occupational social class, mother's highest educational qualification, economic level of the neighborhood, telephone in the home, and ownership/rental of home. Highly correlated, redundant variables (e.g., number of siblings) were substituted by more specific mea-

TABLE 1—Post-Infancy Wheezing to Age 10 by Maternal Smoking at any Time

Reason for Wheezing	Affected Number	Cumulative Incidence		Risk Ratio (.95 CI)
		Smokers	Nonsmokers	
Wheezy bronchitis	625	.074	.052	1.44 (1.24, 1.68)
Asthma	282	.028	.029	.96 (.77, 1.22)
Other reason	632	.064	.067	.96 (.83, 1.11)
Combinations of above	123	.013	.013	1.00 (.71, 1.41)
Total	1662	.180	.162	1.11 (1.02, 1.21)

TABLE 2—Cumulative Incidence of Wheezy Bronchitis in Ages 1–10 by Parental Smoking

Maternal Smoking	Paternal Smoking (Number with wheezy bronchitis)				
	No	Cigarettes	Pipe Cigar	Father Absent	N
No	051 (110)	050 (93)	052 (27)	066 (18)	4815
Yes	066 (59)	076 (238)	064 (20)	083 (44)	4855
Total	055 (169)	067 (331)	056 (47)	077 (62)	9670
Risk Ratio of Maternal effect (.95 CI)	1.29 (.95, 1.75)	1.52 (1.20, 1.92)	1.23 (.70, 2.15)	1.26 (.74, 2.14)	

TABLE 3—Cumulative Incidence Rate of Wheezy Bronchitis at Ages 1–10 by Maternal Smoking Dose

	Average cigarettes per day at birth and 5 years				
	None	1–4	5–14	15–24	>24
Cumulative Incidence Wheezy Bronchitis	.052	.066	.075	.081	.089
Number Children*	4815	925	2450	991	169
Rate Difference	—	.013	.023	.029	.037

Chi-square (trend) = 1.876, p = 0.17

*240 smoking mothers did not smoke at child's birth or age 5 years; 80 had incomplete dose reporting.

TABLE 4—Incidence of Wheezy Bronchitis (WB) by Period of Maternal Smoking

Pregnancy	Child's Age 0–5	Child's Age 5–10	N	# WB	Cumulative Incidence WB	RR WB	95% CI
n	n	n	4815	250	052	—	
y	y	y	3302	260	079	1.52	1.27, 1.82
y	y	n	538	35	065	1.25	0.87, 1.80
y	n	y	307	20	065	1.25	0.79, 2.00
y	n	n	164	11	067	1.29	0.70, 2.39
n	y	y	107	12	112	2.16	1.19, 3.93
n	y	n	197	15	076	1.46	0.86, 2.50
n	n	y	240	10	042	0.80	0.43, 1.51

y = Smoked during pregnancy, or at least three years during interval
n = Did not smoke in pregnancy, or less than three years during interval
RR = Risk ratio

sures (e.g., crowding). Variables unrelated to smoking or wheezing were not included.

The interaction between maternal smoking and the following variables were estimated by both bivariate and multivariate analyses: sex, maternal depression, child allergy, parental allergy, neonatal respiratory distress, breast-feeding, gas cooking, gas heating, crowding, social status, or paternal smoking. The effect of maternal smoking on wheezy bronchitis showed little difference among the levels of any of these factors. Thus, the final multiple logistic models excluded interaction terms.

As Table 5 shows, the dose relationship with maternal smoking appeared to plateau at 15–24 cigarettes per day when all potential confounders were controlled. Male sex, low social status, family allergy, and bedroom dampness, when controlled for maternal smoking, were related to wheezy bronchitis.

Several variables were examined as potential mediators of the smoking effects by adding each separately to the full analytic model, including neonatal respiratory problems, child allergy, and parental respiratory symptoms. These mediating effects were strongest among children whose mothers smoked over 24 cigarettes per day: the maternal smoking effect was reduced (from 1.49 to 1.32) when maternal respiratory symptoms were controlled; parental symptoms exerted no such effect, despite the stronger independent contribution of fathers' report of cough or phlegm to the child's symptoms. A smaller reduction in effect (to 1.36) was evident when maternal depression was included in the analysis. An increase in the effect (to 1.61) was seen when child allergy was held constant. Neonatal respiratory problems showed no mediating effect.

If mothers smoking over four cigarettes per day had not smoked, we estimate that 13.8 per cent of the total burden of wheezy bronchitis in this population would not have occurred.

TABLE 5—Child's Wheezy Bronchitis by Age 10 Years and Maternal Smoking: Multiple Logistic Regression

# cig day* (N)	Odds ratios vs nondiseased children (95% CI) controlled for covariates**
None	1.00
1–4 (925)	1.27 (.95–1.70)
5–14 (2450)	1.43 (1.16–1.76)
15–24 (991)	1.49 (1.13–1.97)
>24 (169)	1.49 (.85–2.63)
*Average of amount between birth and age five	
**Covariates in model, multivariate odds ratio (95% CI):	
Paternal cigarettes	1.05 (.85, 1.29)
Paternal pipe cigar	1.05 (.75, 1.46)
Father absent	1.08 (.81, 1.43)
Male child	1.39 (1.16, 1.64)
Family allergy	1.48 (1.22, 1.78)
Crowding	0.91 (.71, 1.17)
Bedroom dampness	1.67 (1.31, 2.14)
Not breast fed	1.22 (.97, 1.53)
Gas cooking/heating	1.03 (.86, 1.23)
Social status factor	0.90 (.81, .99)

Discussion

In this large, representative national cohort of British children followed from birth, 18 per cent were reported to wheeze by 10 years of age. A subgroup of these children (those whose symptoms were attributed by their mothers to wheezy bronchitis) were at increased risk of post-infancy wheezing when their mothers were smokers. Paternal smoking did not contribute independently to the risk of wheezy bronchitis, perhaps because young children spent more time with their mothers. Recent changes in parental work and child care patterns may alter these relationships.

Is tobacco smoke directly toxic to children's lungs, or are its effects attributable to other factors? Parents who smoke have more respiratory symptoms themselves and may transmit infections to their children.[10,26,27] Additionally, children may imitate the respiratory symptoms of their parents. We found that report of maternal cough or phlegm resulted in some attenuation of the relationship of maternal smoking with the child's wheezy bronchitis; paternal symptoms showed no such relationship. This finding is consistent with mediation of some of the effect of maternal smoking on the child's wheezing by maternal respiratory symptoms; it also is consistent with maternal smoking jointly causing maternal symptoms and the child's wheezy bronchitis, or with more frequent reports of wheezy bronchitis in children of smoking mothers with respiratory symptoms. A substantial portion of childhood wheezy bronchitis was unexplained by these mediating factors, either due to actual toxicity of sidestream tobacco smoke or unexamined mediating factors.

Passive smoking was related only to wheezy bronchitis and not to reported asthma or wheezing for other reasons. We do not know what clinical characteristics led to the perceived diagnosis of wheezy bronchitis in these children. Several investigators have suggested that asthma and wheezy bronchitis are clinically and pathologically indistinguishable, and that they are both manifestations of reactive airway disease.[37,38] The distinct patterns of association of different categories of wheezing with parental smoking found in this cohort may represent true differences between various forms of wheezing.[39] Alternatively, children exposed to tobacco smoke may have been labeled disproportionately as having wheezy bronchitis by health care providers because of earlier age of onset than expected for asthma, relative absence of other atopic symptoms, or awareness of early published reports of an association between passive smoking and bronchitis in childhood.

Reliance on parent report of smoking and retrospective report of childhood wheezing may add uncertainty to our findings.[8,24] However, the causal relationship between pas-

sive smoking and childhood reactive airway disease is supported here by consistency with previous investigations, temporal coherence of the sequence of exposure and symptoms, a dose-response gradient, and biologic plausibility.[40] Our data do not reveal a specific period of vulnerability to the effects of passive smoking. Further follow-up of this cohort is in progress and may add to our understanding of the effects of passive and active smoking on the natural history of reactive airway disease.

We estimate a 49 per cent increase in the prevalence of childhood wheezy bronchitis by age 10 when mothers smoke more than 14 cigarettes per day, when compared with children of nonsmoking mothers. Perhaps as much as 14 per cent of wheezy bronchitis in this population can be attributed to maternal smoking of over four cigarettes per day. These findings add evidence of the public health hazards of cigarette smoke to those involuntarily exposed.

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Nicotine Intake in Young Smokers: Longitudinal Study of Saliva Cotinine Concentrations

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Abstract: Smoking habits and smoke intake were studied over three consecutive years in 197 girls, initially aged 11 to 14 years. Saliva cotinine concentrations in girls who were smokers throughout the three years increased over each year of the study, the greatest increase occurring during movement from occasional to daily smoking. Cigarette consumption also increased over the two years. For a group of continuing daily smokers (n = 23), inhalation of smoke per cigarette as indexed by the ratio of cotinine concentration to average daily cigarette consumption did not change over time. Cotinine concentrations in 16 girls who were smoking on a daily basis within a year of starting to smoke suggested the early development of inhalation. Our findings suggest that young smokers learn to inhale cigarette smoke very early in their smoking careers, that further increases in smoke intake mainly reflect increased cigarette consumption, and that the pharmacological effects of nicotine are already important in reinforcing their smoking. (*Am J Public Health* 1989; 79:172–175.)

Introduction

Nicotine is known to play an important role in smoking addiction in adults,[1] but has been little studied in the early stages of smoking in children.

Questionnaire measures such as cigarette consumption give only a rough estimate of nicotine intake[2] and need to be supplemented by biochemical measures. Nicotine has a relatively short half-life (approximately two hours) in the body so that its plasma concentration is primarily a measure of the last few cigarettes smoked.[3] Cotinine, a major metabolite of nicotine, has a longer half-life (approximately 16 hours[4]) and its concentration in plasma and saliva has been used as a marker of nicotine intake in adults.[3] Quantitative estimation of smoke intake in the early stages of smoking requires objective measures which are sensitive even to intermittent smoking, and it has recently been shown that saliva cotinine provides such a measure.[5,6]

Analysis of cross-sectional data has suggested an early development of smoke intake in young smokers. Inhalation as reflected in cotinine concentrations and expired air-carbon monoxide values was already established in teenage smokers and showed little tendency to increase with age,[5,7] suggesting that smoke inhalation may not increase with the number of years of smoking.

This paper reports on a longitudinal study of smoking which included measurement of saliva cotinine and was carried out over three consecutive years from 1985 to 1987 among the pupils attending a girls' comprehensive school in inner London. The study focuses on girls who were smoking at the outset or who took up smoking during the course of the study. Cross-sectional results for the whole school in 1985 have been reported elsewhere.[5,7]

Method

Smoking surveys were scheduled as a curriculum activity for each class between January and March in 1985, 1986, and 1987. Testing was carried out usually in the late morning or early afternoon, and one further attempt was made to test those who had been absent. The pupils did not know in advance that they would be tested, although they were aware that the program was under way. On each occasion, the girls completed a smoking questionnaire and provided saliva samples for cotinine analysis. Before filling in the questionnaires, they were told that the biochemical measure could detect smoking.

Saliva was collected using cotton dental rolls which were placed in the mouth between cheek and gum during the 20 minutes it took to complete the questionnaires. The saliva was subsequently extracted and frozen until analysis for cotinine by gas chromatography.[8]

Repeated measures analyses of variance were carried out using the SPSSX Manova routine to investigate trends over time in cotinine concentrations and cigarette consumption.

Results

Three hundred seventy-four girls (91 per cent of the number on roll in the first, second, or third years) were present and surveyed in 1985. Their ages ranged from 11 to 14 years (mean 13.0, SD=0.8). A cohort of 197 girls (53 per cent of those surveyed in 1985, 60 per cent of those still at the school) had complete data at all three points of the survey (Table 1).

The smoking prevalence (defined as any current smoking) of this cohort increased from 28 per cent in 1985 to 36 per cent in 1987 (Figure 1). Of the 54 self-reported smokers in 1985, seven reported having quit by 1987. Of those smoking in 1985, there were 27 (50 per cent) daily smokers of whom only one had quit by 1987.

Continuing Smokers

Forty-three girls were self-reported smokers on all three occasions. Their cotinine concentrations increased linearly over time (a mean increase of 52.4 ng/ml per year, Table 2). There was no evidence of a quadratic trend, indicating no slowing down of the increase in the second year.

Eighteen (42 per cent) of these 43 smokers classified themselves as occasional smokers in 1985. By 1986, 12 (67 per cent) of these had progressed to daily smoking and were still daily smokers in 1987 (subsample 1, Table 2). Within this

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TABLE 1—Number (%) of Adolescent Girls Eligible for Follow-up Participating in Each of Three Surveys

	1985	1986	1987	Follow-up Cohort
No. (%) present and surveyed	374 (100)	300 (80)[a]	278 (74)	250[b] (67)
No. (%) with complete questionnaire and cotinine data	322 (86)	285 (76)	259 (69)	197[c] (53)

a) Percentages are of those girls present and surveyed in 1985.
b) Losses to follow-up were due to illness or truancy (n = 76) or having moved to another school (n = 48).
c) Missing data were due to incomplete questionnaires (n = 6), refusal to give saliva (n = 1), and those saliva samples discarded due to contamination or insufficient volumes obtained (n = 46). The latter were divided equally between self-reported smokers and non-smokers, the majority (n = 30) occurring in 1985 due to the dental rolls being soaked for an inadequate length of time. Ensuring that they were kept in the mouth for a minimum of 20 minutes largely removed this problem in the later surveys.

group there was a clear linear trend in cotinine concentrations and a suggestion of a quadratic effect, the latter indicating that the increase was greater during the first year of the study when they moved from occasional to daily smoking.

Within a group of 23 girls who were self-reported daily smokers on all three occasions (subsample 2, Table 2) cotinine concentrations showed no quadratic change but increased linearly over the course of the study.

Substantial differences were apparent in the changes in cotinine concentrations between these two subsamples from 1985 to 1986 (difference = 79.9; 95 per cent CI = 7.5,152.3) but not from 1986 to 1987 (difference = 13.0; 95 per cent CI = −54.7,80.7) by which time both groups were smoking on a daily basis.

Usual weekly cigarette consumption increased linearly over time for all smokers. For the subsample of 1985 occasional smokers who were daily smokers in 1986 and 1987, a quadratic trend was also apparent (Table 2).

A question of interest is whether there was any change in smoke intake per cigarette over the two years. A direct measure of this was not taken but it can be indexed approximately by the ratio of cotinine concentration to usual daily cigarette consumption. A repeated measures analysis of variance indicated no change in this index over time for the 23 continuing daily smokers (data available on request to authors). For occasional smokers, the infrequent nature of their smoking, and the ambiguity arising from passive smoking cotinine levels, would make a similar analysis unreliable.

Girls Taking up Smoking During the Study

Thirty-five (25 per cent) of the 1985 non-smokers took up smoking over the following two years (23 in 1986, 12 in 1987). Their mean cotinine concentration after less than one year of smoking was 57.2 ng/ml (SD = 85.1) and their mean weekly cigarette consumption 24.1 (SD = 69.9). Sixteen (64 per cent) of these girls had taken up *daily* smoking within a year and had an end of year mean cotinine concentration of 87.8 ng/ml (SD = 92.8) and daily cigarette consumption of 4.6 (SD = 3.9).

Of the 23 girls who had taken up smoking in 1986, 12 (52 per cent) were still smoking one year later; their cotinine concentrations increased substantially over the next year (Table 2). Nine (75 per cent) of these girls were smoking on a daily basis by 1987.

Discussion

To our knowledge, the stability of cotinine or nicotine concentrations over a period of years has not yet been determined in adults. However, it is known that steady-state blood nicotine concentrations are maintained from one day to the next[10] and this stability would probably remain over a longer period of time. Two recent surveys of adult smokers give average saliva cotinine concentrations of 350 ng/ml for cigarette consumption averaging 25 cigarettes a day.[11,12] This stability contrasts with the situation found in this group of

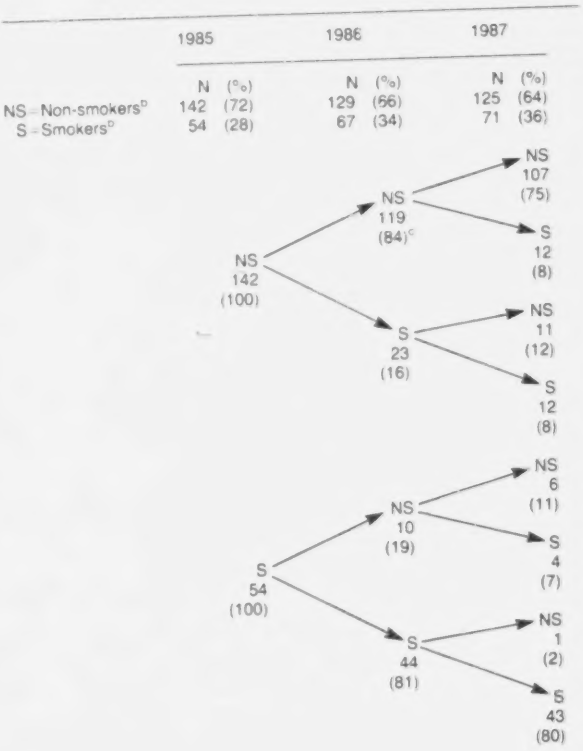

FIGURE 1—Smoking Behavior of the Follow-up Cohort (n=196[a])

a) One self-reported non-smoker in 1985 had a cotinine concentration above a smoking/non-smoking cut-off[4] and was excluded from further analyses.

b) The measure of smoking behavior was a version of Bewley's[9] scale: 1="I have never smoked a cigarette, not even a puff"; 2="I have only ever tried smoking once or twice"; 3="I used to smoke sometimes but I don't now"; 4="I smoke sometimes but don't smoke as much as 1 cigarette a week"; 5="I usually smoke between 1 and 6 cigarettes a week"; 6="I usually smoke more than 6 cigarettes a week." Non-smokers were defined as those who classified themselves in categories 1 to 3, smokers in categories 4 to 6.

(c) Percentages (in parentheses) are of 1985 Non-smokers or Smokers.

TABLE 2—Mean (SD) Cotinine Concentrations and Cigarette Consumption Measures for Continuing Smokers and New Smokers.

				Trend Coefficients (95% CI)	
	1985	1986	1987	Linear Component	Quadratic Component
Smokers at all Three Points of the Survey (n=43)					
Cotinine concentrations	103.6 (89.9)	158.7 (109.7)	208.5 (120.3)	52.4 (35.6,69.2)	−2.7 (−28.9,23.5)
Weekly cigarette consumption	30.9 (44.9)	39.3 (30.7)	55.7 (36.9)	14.4[b] (8.2,20.2)	4.1 (−2.9,11.1)
Subsample 1: 1985 Occasional[c] Smokers, Daily in 1986 and 1987 (n=12)[d]					
Cotinine concentrations	35.7 (49.1)	157.4 (101.7)	206.7 (146.8)	85.5 (42.3,128.7)	−24.3[b] (−62.1,3.8)
Weekly cigarette consumption	2.3 (1.5)	32.5 (18.3)	40.2 (21.0)	19.0 (12.2,25.7)	−11.3 (−21.0,−1.6)
Subsample 2: Daily Smokers in all Three Years (n=23)[e]					
Cotinine concentrations	159.5 (78.6)	201.4 (99.2)	237.7 (94.0)	39.1 (23.8,54.4)	−2.8 (−42.5,37.0)
Daily cigarette consumption	7.6 (7.4)	9.7 (6.0)	11.2 (4.7)	2.2[b] (1.3,3.0)	−0.3[b] (−1.0,1.4)
Non-smokers in 1985, Smokers in 1986 and 1987 (n=12)					
Cotinine concentrations	5.3 (3.7)	57.9 (75.6)	149.4 (111.1)	91.5 (12.7,170.2)	
Weekly cigarette consumption	—	14.3 (18.5)	50.5 (39.8)	36.3 (14.7,57.8)	

a) Weekly cigarette consumption was defined by the following questions: "How many cigarettes have you smoked in the last week" (1986); "How many cigarettes have you smoked in the last seven days" (1987). The 1985 measure was derived from self-reported daily cigarette consumption and the smoking status question from that year.[5]

b) Due to a skewed distribution in these components, log transformations were used.

c) Daily and occasional smokers were differentiated by the question: "If you smoke, do you usually smoke every day?"

d) Of the remaining eight smokers, six occasional smokers continued to smoke on an occasional basis in 1986 (of whom four had taken up daily smoking by 1987) and two of the daily smokers reverted to occasional smoking during the course of the study (one of whom had resumed daily smoking by 1987).

e) This measure was only available for the self-reported daily smokers.

adolescent girls where cotinine concentrations increased substantially over time. Although the greatest rise in cotinine concentrations occurred during movement from occasional to daily smoking, smoke intake continued to increase albeit less rapidly, on attaining a regular daily smoking pattern. Cotinine concentrations even at the beginning of the study were substantial, with approximately half the average adult concentration being observed in the 11 to 14 year old daily smokers. Two years later the same group had levels more than two-thirds those in the adult surveys. This is in spite of the fact that these young smokers were at school and subject to considerable restrictions on their smoking behavior and probable financial limitations on the number of cigarettes they could smoke.

Hence, these girls were receiving substantial doses of nicotine from a very early stage in their smoking careers. This suggests that pharmacological effects of nicotine were already important in perpetuating their smoking. Other analyses of the 1985 data indicate an early development of pharmacological addiction: young smokers reported experiencing calming effects when smoking and withdrawal effects during attempts to give up which were similar to those experienced by adult smokers.[13,14]

Since cigarette consumption in our sample also increased over the two years, it seems likely that the increase in cotinine concentrations in our sample was due in part to an increase in cigarette consumption. Indeed, an index of smoke intake per cigarette computed for the continuing daily smokers showed no tendency to increase over the two years. An early development of inhalation is also suggested when examining the cotinine concentrations for the 16 adolescents who had been smoking for less than one year yet were already daily smokers. Nevertheless, changes in nicotine metabolism, cotinine half-life or simply a lack of precision in the cigarette consumption measures in this sample of adolescents could disguise the true picture and growing adolescents would have to increase their smoke intake over time to counteract any concomitant increases in body weight. Therefore, we cannot rule out the possibility of increases in the amount of smoke inhaled per cigarette during these years.

Although this sample was made up entirely of girls, there is no *a priori* reason to expect the findings would not also apply to boys. Attrition of the follow-up cohort due to sickness, truancy and leaving the school may mean that there was some selection bias in our sample. It is also possible that the study itself may have influenced smoking behavior. However, the likely direction of such effects would mean that our findings would be conservative.

Our results indicate that nicotine is already important in maintaining the smoking habit in adolescents. Obviously, interventions are needed well before this stage is reached and preferably before the onset of smoking, which in many cases would appear to be before arrival at secondary school.[15]

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Surgeon General C. Everett Koop Receives APHA Award for Excellence

Editor's Note: At its 116th annual meeting in Boston, November 13–17, 1988, the American Public Health Association conferred its 1988 Award for Excellence on C. Everett Koop, MD, United States Surgeon General, in recognition of his "exceptional contributions to the improvement of the health of the people." During an Honors Assembly held November 14, the Association recognized the "meritorious contributions" made by Dr. Koop during his seven years as Surgeon General. Portions of the citation read in honor of Dr. Koop dealt particularly with his leadership role in relation to the tobacco issue. This issue of the Journal focuses on the subject of tobacco, smoking, and smoking cessation in honor of the 25th anniversary of the Surgeon General's first report on smoking and health. Excerpts from the citation honoring Dr. Koop last November are printed below:

Excerpts from APHA Citation for Surgeon General Koop

". . . .Dr. Koop has maximized the symbolic authority of the Surgeon General's office . . . Through a fortuitous conjunction of his own substantial leadership qualities with the extraordinary health challenges of our day has come a dramatic and lasting impact upon our society and the health of our people. . . .

"As the nation's most recognized spokesperson for public health, he has used the high visibility of his office with great effectiveness to mold public opinion and to stimulate societal action for the prevention of disease and the promotion of health, focusing particularly upon two of the most pressing health issues of our day—tobacco and AIDS. It is especially for these achievements we cite Surgeon General Koop.

"Soon after becoming Surgeon General, he surprised some by continuing to speak out forcefully and frequently concerning the adverse health effects of smoking. He became the first federal official to speak out explicitly and authoritatively for a 'smoke-free society.' With increasing public attention, he personally promoted the now traditional annual Surgeon General's reports on smoking and health. In one, he openly documented the compelling evidence against passive smoking. In another, he boldly and soundly exposed the addictive nature of tobacco. Yet another targeted the expanding consequences of smoking on women's ill health. He unabashedly assaulted the commercialization of disease and death by the tobacco industry, and endorsed the desirability of a ban on the advertising of tobacco products.

. . .

"The American Public Health Association is pleased to present its 1988 Award for Excellence to Surgeon General C. Everett Koop in recognition of his exceptional contributions to the improvement of the health of the people."

A Survey of Smoking and Quitting Patterns among Black Americans

Carole Tracy Orleans, PhD, Victor J. Schoenbach, PhD, Mary Anne Salmon, PhD, Victor J. Strecher, PhD, William Kalsbeek, PhD, Dana Quade, PhD, Edward F. Brooks, DrPH, T. Robert Konrad, PhD, Charles Blackmon, MBA, and Charles D. Watts, MD

Abstract: A sample of adult Black policyholders of the nation's largest Black-owned life insurance company was surveyed in 1986 to add to limited data on smoking and quitting patterns among Black Americans, and to provide direction for cessation initiatives targeted to Black smokers. Forty per cent of 2,958 age-eligible policyholders for whom current addresses were available returned a completed questionnaire. Population estimates for smoking status agree closely with national estimates for Blacks age 21–60 years: 50 per cent never-smokers; 36 per cent current smokers; 14 per cent ex-smokers. Current and ex-smokers reported a modal low-rate/high nicotine menthol smoking pattern. Current smokers reported a mean of 3.8 serious quit attempts, a strong desire and intention to quit smoking, and limited past use of effective quit smoking treatments and self-help resources. Correlates of motivation to quit smoking were similar to those found among smokers in the general population, including smoking-related illnesses and medical advice to quit smoking, previous quit attempts, beliefs in smoking-related health harms/quitting benefits, and expected social support for quitting. Methodological limitations and implications for the design of needed Black-focused quit smoking initiatives are discussed. (*Am J Public Health* 1989; 79:176–181.)

Introduction

Promoting smoking cessation among Black Americans has been declared a public health priority.[1] Blacks, who account for 12 per cent of the US population, constitute the nation's largest minority group and have the highest smoking rate: 34.8 per cent of all Black Americans smoke compared to 29.7 per cent of non-Hispanic White and 25.7 per cent of Hispanic adults.[2] Among Black Americans, as for the US population as a whole, smoking rates are highest, and quit rates are lowest, within lower income and blue-collar subgroups.[3,4] Blacks also suffer the nation's highest rates of mortality and morbidity from smoking-related cardiovascular diseases and cancer—including coronary heart disease and lung cancer.[4–6] Moreover, smoking compounds other serious health risks within the US Black population, including a disproportionate incidence of infant mortality and low birthweight, hypertension, and diabetes.[6]

To date, relatively little research has been done to clarify smoking and quitting patterns and determinants among Black Americans, or to test quit smoking treatments in Black populations.[3,7,8] This paper presents results of a survey on the smoking and quitting patterns of lower and lower-middle income Black Americans. The survey was conducted to aid in the design of needed Black-focused quit smoking initiatives and to prepare for a smoking cessation intervention trial among policyholders of the North Carolina Mutual Life Insurance Company, the nation's largest Black-owned life insurance company.

Methods

Study Population

The target population for the survey consisted of North Carolina Mutual (NCM) Life Insurance Company policyholders aged 21–60 with active individual life insurance policies as of December 1985 and residing in 12 Southern, Mid-Atlantic, and Midwestern states and the District of Columbia where the company maintains active sales offices. All individual policyholders are Black and from predominantly low and lower-middle income groups. A stratified simple random sample for 4,000 NCM policyholders aged 20–61 was drawn from an unduplicated computer file of approximately 178,000 holders of individual life insurance policies in 37 geographic sales districts. To improve the statistical quality of survey estimates, selection of the sample was stratified by sex, age, and whether or not the policyholder had an industrial-type policy. Industrial policies (69 per cent of policyholders) have weekly or monthly agent collection of premium payments ("home service") and face values mostly under $1,000; ordinary policies are of either "home service" or mailed premium types and have face values of at least $1,000. Males were slightly oversampled to achieve a more nearly sex-balanced sample than the original policyholder sampling frame (57 per cent female).

Data Collection

The survey was conducted from April-September 1986. A variety of organizational factors and research objectives dictated the design of data collection procedures. At the time of the survey, the NCM home office did not maintain current addresses or telephone numbers for industrial policies. Its approximately 500 field agents, who see most policyholders on a regular basis, were regarded as the key to obtaining access to and cooperation of policyholders. Furthermore, since agents were to play a major role in a smoking cessation trial to follow this survey, involving them in data collection was essential to building a collaborative research relationship and assessing their research capabilities.

Agents were asked to:

- supply up-to-date mailing addresses and telephone numbers for all of their policyholders selected into the sample;
- notify subjects about the survey before mailing the survey questionnaires or deliver them in person;
- make a follow-up call or visit several days later to encourage cooperation and answer questions; and
- visit nonrespondents in person to provide further encouragement or assistance if needed.

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District-based group training sessions instructed agents in how to introduce the survey and encourage participation, and included both videotaped modeling of agent contacts and printed guidelines emphasizing confidentiality of responses. To minimize policyholder concerns about confidentiality, a cover letter attached to the survey explained the purpose of the study with assurances that participants would be identified only by code number, that completed surveys would be delivered in a sealed envelope to the research team at the University of North Carolina, and that individual results would neither be reported to NCM nor ever affect insurance rates or coverage. Cooperation was encouraged through modest lottery-style financial incentives for agents and policyholders, with primary reliance for agent compliance on the company's management structure. For reasons of policyholder and agent relations, NCM management discouraged direct contact with nonrespondents.

The "Smoking and Health Survey" requested sociodemographic information from all respondents (one page) and included separate additional sections for smokers (four pages) and ex-smokers (three pages).* Current smokers were asked about smoking rate, brand (brand name and Uniform Product Code number), nicotine dependence,[10] past attempts to quit or cut down, current desire to quit, beliefs about smoking health harms and quitting benefits, presence of smoking-related illnesses, and interest in self-help materials. Ex-smokers were asked about past smoking and quitting history and reasons for quitting. Standard smoking history items from the National Cancer Institute's self-help smoking cessation trials were used whenever possible.[11] The questionnaire was carefully pilot-tested for readability and ease of completion.

Results

Agents provided usable mailing addresses for 2,958 sample policyholders and reported an additional 249 as having lapsed policies. Withdrawing these 249 inactive policyholders, 107 potential subjects in a district selected for pretesting and 25 served by district offices no longer active yielded a potential sample of 3,619 subjects. Completed surveys were received from 1,237. Eliminating 74 respondents (6 per cent) ineligible on the basis of age left 1,163 eligible respondents. Counting as nonrespondents both those for whom agents did not supply an address and those who failed to return a completed questionnaire yields a 32 per cent response rate (1,163/3,619). Restricting the denominator to policyholders for whom agents supplied mailing addresses and correcting for estimated age-eligibility yields a 40 per cent participation rate (1,163/2,884).

Response rates for the 35 study sales districts ranged from 7 per cent to 62 per cent (median 38 per cent). Although there were regional patterns (generally higher rates in the South [15–62 per cent] and lower rates in the Midwest [10–51 per cent] and Mid-Atlantic [7–47 per cent]), there were marked differences within the same state or even the same city. Apparently valid addresses were obtained for 80 per cent or more of subjects in 25 (71 per cent) of the districts; seven districts provided addresses for only 45–69 per cent of subjects. Differences in district response rates appeared to be due to differences in access to policyholder populations and in agent and management enthusiasm for the survey. No relationships between response rates (as a percentage of the sample or as a percentage of apparently valid addresses) and prevalence of smokers, ex-smokers, and never-smokers were apparent in scatter plots.

Characteristics of respondents were compared with those of the total eligible sample in order to identify any systematic nonresponse patterns. Among the 16 sampling strata formed by sex, age, and type of policyholder, response rates varied from 22 per cent to 43 per cent. Response rates were 26 per cent (about 10 percentage points) lower for industrial policyholders, due to the unavailability of 26 per cent of their addresses. Rates for males were 11 per cent (4 percentage points) lower than for females, in both industrial and ordinary categories. Industrial policyholders, males, and industrial male policyholders age 40 years and younger were slightly underrepresented in the (unweighted) respondent sample (63 per cent vs 69 per cent, 46 per cent vs 49 per cent, 15 per cent vs 20 per cent, respectively, for respondents vs eligible sample).

To help compensate for nonresponse bias and to correct for the disproportion resulting from intentional oversampling of males, all estimates presented in this paper were weighted to the age, sex, and policy type distribution of the policyholder population. This was accomplished by means of a standard adjustment whereby, in effect, the mean of similar respondents was used as a proxy for each nonrespondent.[12] Estimates and variances were calculated using a computer program designed to properly account for the complexity of the sampling design.[13]

Smoking Status and Sociodemographic Data

Based on data from the 1,163 respondents, it was estimated that 50 per cent of NCM's policyholders were never-smokers, 36 per cent were current smokers, and 14 per cent were ex-smokers (see Table 1). The majority of policyholders were women, with a mean age of 39 years. Fewer than half were married. Most had a high school education or less, and most were employed full-time. Respondents reported a mean of 2.5 ambulatory medical visits in the past year, and 84 per cent described their health as "good" or "excellent."

Two-sample t-tests (two-tailed) and chi-squares showed expected relationships of smoking status to age, sex, marital status, education, employment and health.** Never smokers were predominantly women. Ex-smokers were older than current or never smokers. Fewer current smokers than ex-smokers or never smokers were married, employed full-time, or in excellent/good health. Both current smokers and ex-smokers were less likely to have completed high school than never smokers. One-third of current smokers, but half of ex-smokers, reported above normal weight for height and sex, and the mean Quetelet (weight/height) index was significantly higher for ex-smokers than for current smokers.[14]

Current Smokers' Smoking and Quitting Patterns

Current smokers began smoking at a mean of 17.5 years and had smoked for a mean of 19.4 years (Table 2). They were predominantly low-rate smokers, reporting an average of 14

*A detailed report of data collection and agent training methods and a copy of the survey instrument can be obtained from Mary Anne Salmon, PhD, Health Services Research Center, University of North Carolina, Chapel Hill, NC 27599-7490.

**More detailed documentation of sampling, weighting, estimation and analysis procedures can be obtained from William Kalsbeek, PhD, Department of Biostatistics, School of Public Health, University of North Carolina, Chapel Hill, NC 27599.

TABLE 1—Sociodemographic Characteristics of Individual North Carolina Mutual Policyholders by Smoking Status, 1986

Characteristics	Current Smokers (N = 399)	Ex-Smokers (N = 175)	Never Smokers (N = 589)	Total (N = 1163)
% Smoking Status	35.8 (1.7)	14.2 (1.2)	49.9 (1.7)	100.0
Mean age	37.2 (0.6)	45.1 (1.0)	38.5 (0.5)	39.0 (0.3)
% Female	49.5 (2.7)	47.4 (4.2)	69.8 (1.9)	59.3 (0.6)
Education				
% < High School	32.8 (2.8)	33.8 (4.1)	20.1 (1.8)	26.5 (1.5)
% High School	40.7 (3.1)	30.5 (4.3)	36.4 (2.3)	37.1 (1.7)
% > High School	26.5 (2.7)	35.7 (4.3)	43.5 (2.4)	36.4 (1.7)
Employment				
% Employed Full-time	49.8 (3.1)	69.4 (4.0)	65.9 (2.3)	60.6 (1.7)
% Married	27.8 (2.6)	52.0 (4.5)	43.2 (2.4)	38.9 (1.6)
Health Status				
% Excellent Good	77.5 (2.6)	84.5 (3.7)	88.7 (1.6)	84.2 (1.3)
% Fair Poor	22.5 (2.6)	15.5 (3.7)	11.3 (1.6)	15.8 (1.3)
Mean medical visits past year	2.4 (0.2)	3.0 (0.3)	2.4 (0.3)	2.5 (1.7)
Relative weight of smokers and ex-smokers				
% Normal or Low	66.7 (2.8)	49.0 (4.5)		
% Moderate Overweight	10.3 (1.8)	13.4 (2.8)		
% Severe Overweight	23.1 (2.6)	37.6 (4.4)		
Mean Quetelet Index	29.5 (0.5)	32.1 (1.1)		

Estimates are weighted to compensate for stratification and nonresponse. Standard errors are given in parentheses. No variable missing in >10 per cent of cases.

cigarettes a day. Fifty-two per cent smoked 10 or fewer cigarettes a day, and only 12 per cent smoked 21 or more cigarettes a day. However, most smoked high nicotine brands and most smoked within 30 minutes after waking, a crude index of high nicotine dependency.[10] Almost three-fourths smoked menthol brands; three menthol brands that dominate among Black smokers[15] (Kool, Newport, and Salem) accounted for 55 per cent of brands reported. Only 11 per cent of current smokers used tobacco products other than cigarettes.

Smokers in this population were estimated to have a mean of 3.8 lifetime quit attempts, relying mainly on self-change strategies and will power. Almost 80 per cent had tried once or more to quit, but fewer than 5 per cent had ever tried a formal group or individual treatment. Likewise, more had tried ineffective drugstore aids (14 per cent) than either self-quitting guides (9 per cent) or nicotine gum (7 per cent).

About nine in 10 current smokers had cut down their smoking in one or more ways during the past year. Cutting down on the number of cigarettes smoked and cutting down on smoking around nonsmokers were more common than either inhaling less of each cigarette or switching to lower nicotine brands.

Approximately one-fourth of current smokers had suffered from a smoking-related chronic illness in the past year (hypertension, heart disease, diabetes, bronchitis, emphysema, asthma and/or lung cancer), and over one-third had received personal medical advice to quit in the past year. Smokers generally expressed a moderately strong desire to quit smoking and strong beliefs in smoking health harms and quitting benefits, but they expected only moderate social support for their quitting efforts. Two-thirds were "seriously thinking about quitting smoking in the next year," placing them in the "contemplation stage" as defined by Prochaska and DiClemente.[16] Nine in 10 current smokers were interested in receiving a free self-quitting guide from NCM; 64 per cent were "very" interested.

To identify correlates of strong quitting motivation, the 47 per cent of current smokers with maximum "desire to quit" ratings on an 11-point scale were compared with the remaining 53 per cent on sociodemographic and smoking history variables. As Table 3 shows, the following variables were significantly associated with a strong desire to quit: female gender; lower nicotine cigarette brand; greater number of cut down strategies and previous quit attempts and methods; smoking-related illness or medical advice to quit smoking in the past year; stronger beliefs in smoking health harms and quitting benefits; and greater social support expected for quitting.

Ex-Smokers' Smoking and Quitting Patterns

Ex-smokers and current smokers were similar in regard to mean age of smoking onset (18 years) and mean smoking history (17.5 years) (see Table 2). Ex-smokers were older than current smokers and had been off cigarettes for an average of 10 years; only 9 per cent presently used other tobacco products. Like current smokers, ex-smokers had a relatively low daily smoking rate before quitting (15.5 cigarettes/day). Fewer ex-smokers than current smokers had routinely smoked within 30 minutes after waking.

As shown in Table 4, ex-smokers cited health concerns as foremost reasons for quitting: the desire to overcome present smoking-related symptoms was most often noted, followed by the desire to avoid future illness. Self-mastery motivation and quitting to set a better example for children or others were the next most common reasons. A variety of other motivations—economic, religious, cosmetic, social—were also endorsed.

We also asked ex-smokers to recall the quitting and cut down methods they tried (Table 2). Fewer ex-smokers than current smokers reported cut-down strategies: this difference was most pronounced for "cutting down around others who did not smoke," likely reflecting the major shift over the past decade in smoking norms and in knowledge and attitudes about second-hand smoke. Comparably small proportions of ex-smokers and current smokers had used formal treatments.

TABLE 2—Smoking and Quitting History for Current and Ex-Smokers

Variables	Current Smokers (N = 399)	Ex-Smokers (N = 175)
Mean age started smoking	17.5 (0.3)	18.1 (0.3)
Mean age quit (ex-smokers only)		35.5 (0.8)
Mean years quit (ex-smokers only)		9.5 (0.8)
Mean years smoked	19.4 (0.6)	17.5 (1.1)
Mean cigarettes per day	13.9 (0.6)	15.5 (1.1)
% smoking		
≤10 cigarettes day	51.6 (3.1)	50.9 (4.5)
11–20 cigarettes day	36.6 (3.0)	30.2 (4.0)
21–30 cigarettes day	6.9 (1.8)	10.7 (1.5)
31+ cigarettes day	4.9 (1.5)	8.2 (2.2)
% Smoke <30 minutes after waking	63.8 (2.9)	39.4 (4.2)
Mean mgs nicotine per cigarette	0.98 (.02)	
% High nicotine brand (≥1.0 mg)	63.1 (3.1)	
% Medium nicotine brand (.7–1.0 mg)	28.3 (2.5)	
% Low nicotine brand (<.7 mg)	8.6 (1.8)	
% Smoking menthol brands (smokers only)	73.3 (2.6)	
% Other tobacco past month:		
Daily	1.2 (0.5)	3.6 (1.8)
Sometimes	9.8 (1.7)	5.3 (0.6)
Never	89.0 (1.8)	91.1 (2.4)
Mean lifetime quit attempts (smokers only)	3.8 (0.2)	
% ≥1 lifetime quit attempt (smokers only)	79.3 (2.5)	
Mean ways cut down (past year or prequit)	2.0 (0.1)	1.4 (0.1)
% cut number of cigarettes	60.8 (2.9)	50.4 (4.4)
% switched to lower nicotine	26.7 (2.6)	16.2 (3.3)
% cut down around others	62.1 (2.9)	39.6 (4.3)
% inhaled less/less of each cig	52.6 (3.0)	37.7 (4.2)
Mean quitting methods tried (past year or prequit)	1.3 (0.1)	1.5 (0.1)
% Any formal treatment	4.1 (1.3)	2.7 (1.3)
group	1.1 (0.6)	0.4 (0.4)
one-to-one	3.3 (1.2)	2.8 (1.3)
% Books/guides	9.2 (1.8)	3.2 (1.4)
% Nicorette	7.3 (1.5)	1.2 (0.9)
% Any drugstore aid	14.2 (2.1)	3.3 (1.3)
filters	12.3 (2.1)	2.4 (1.1)
lozenges	5.9 (1.5)	1.5 (0.7)
% Prayer or meditation	19.0 (2.4)	21.7 (3.4)
% Will power	71.5 (2.8)	89.3 (3.4)
% With smoking-related illness past year	25.4 (2.4)	
% With medical advice to quit past year	35.0 (2.9)	
Mean belief smoking harms health (Scale 0–10)	8.5 (0.2)	
Mean belief quitting improves health (Scale 0–10)	8.5 (0.2)	
Mean desire to quit (Scale 0–10)	7.2 (0.2)	
Mean expected social support for quitting (Scale 0–10)	6.0 (0.3)	
% Seriously thinking of quitting in next year	67.6 (2.9)	
% Interested in free guide from NCM:		
Very interested	64.1 (3.0)	
Somewhat interested	26.4 (2.7)	
Not interested	9.6 (1.8)	

Estimates are weighted to compensate for stratification and nonresponse. No variable missing in ≥10 per cent of cases, except cigarette brand information, missing in 30 per cent of cases. Standard errors are given in parentheses.

Fewer ex-smokers than current smokers reported trying quit smoking books or guides, drugstore aids, or nicotine gum (not available in the US until 1984). More ex-smokers than current smokers reported relying on will power to succeed.

Discussion

The low response rate is obviously troublesome, but the final respondent sample appears reasonably representative of the eligible sample, and survey data were weighted to help compensate for the slight underrepresentation of industrial policyholders, males, and younger male industrial policyholders. Such weighting would not correct for the possibility that policyholders with a greater interest in smoking as a health issue (and/or whose agents were more vigorous in obtaining addresses and encouraging participation) are overrepresented in our respondent sample. Likewise, despite the strong emphasis on confidentiality of results, respondents may have attempted to present their behavior and intentions in a favorable light, particularly given the survey's identification with their insurance company. These factors would serve to create a "best case" bias, limiting generalizability, although probably not greatly limiting the usefulness of the survey to inform the design of quit smoking materials and programs for smokers interested in quitting.

Findings from this study are not directly generalizable to

TABLE 3—Sociodemographic and Smoking History Correlates of Strong Desire to Quit Smoking among Current Smokers (N = 399)

Correlates	Want to Quit = 10 (Scale 0–10) N = 185	Want to Quit = 0–9 (Scale 0–10) N = 214
Sex (% female)	57.8 (3.7)	42.1 (3.6)
Mean mgs nicotine per cigarette	0.93 (0.3)	1.0 (0.2)
Mean ways cut down (past year or prequit)	2.3 (0.1)	1.9 (0.1)
% cut number of cigarettes	73.2 (3.7)	52.7 (4.2)
% cut down around others	71.6 (3.8)	56.7 (4.3)
Mean number of prior quit attempts	5.1 (0.3)	2.6 (0.3)
Mean quitting methods tried (past year or prequit)	1.5 (0.1)	1.1 (0.1)
% Nicorette®	10.6 (2.6)	4.4 (1.4)
% Prayer/meditation	27.3 (3.9)	11.9 (3.0)
% Will power	84.3 (3.3)	59.9 (4.2)
% With smoking-related illness past year	33.6 (4.0)	19.6 (3.0)
% With medical advice to quit past year	43.8 (4.3)	27.3 (3.8)
Mean belief smoking harms health (Scale 0–10)	9.4 (0.2)	7.7 (0.2)
Mean belief quitting improves health (Scale 0–10)	9.5 (0.1)	7.6 (0.2)
Mean expected support for quitting (Scale 0–10)	7.1 (0.3)	5.0 (0.4)
% Very interested in free guide from NCM	87.9 (3.1)	42.1 (4.2)

Estimates are weighted to compensate for stratification and nonresponse. Standard errors are given in parentheses. No variable missing in ≥10 per cent of cases, except mean mgs. nicotine per cigarette, missing in 30 per cent of cases.

TABLE 4—Reasons for Quitting Endorsed by Ex-Smokers (N = 175)

Reason for Quitting	% Endorsing
To feel better physically	77.3 (3.7)
To prevent future illness	63.4 (4.2)
To take more control over my life	53.5 (4.3)
To set example for children/others	42.7 (4.2)
Don't like smell/stains teeth	33.2 (4.2)
To please or help a loved one	29.7 (3.9)
To save money	28.2 (3.9)
So my children wouldn't breathe cigarette smoke	23.0 (3.6)
Because my religion is against it	19.3 (3.3)
Because my doctor told me to	17.3 (3.6)
Because of immediate health problems	17.1 (3.5)
Because my friends don't smoke	8.5 (2.9)
Because of pregnancy (women only)	7.7 (3.1)

Estimates are weighted to compensate for stratification and nonresponse. Standard errors are given in parentheses. No variable missing data for ≥10 per cent of cases.

the general population of Black Americans. Compared to US Black adults 21–60 years, the weighted NCM respondent sample had a slightly higher percentage of females (59 vs 55 per cent); was somewhat older (38 vs 34 per cent of males were older than 40 years, as were 48 vs 36 per cent of females); and had somewhat greater formal education (26.5 vs 29.5 per cent had not completed high school and 36 vs 30 per cent had post-secondary education). It is reassuring that despite the high nonresponse rate the smoking prevalence (36 per cent) and ex-smoking prevalence (14 per cent) estimated for the policyholder population agree fairly closely with the national data (40 per cent and 15 per cent, respectively), suggesting only a small nonresponse tendency among current and former smokers.

Since this paper focuses on smokers and ex-smokers, the comparison of our sample of policyholders to the US by smoking status is relevant. Among current smokers, policyholders had the same sex and age distributions and slightly higher educational levels as for US Blacks (33 vs 37 per cent had not completed high school; 26.5 vs 23.3 per cent had post-secondary education). Among ex-smokers, the proportion of females in NCM was three percentage points higher than for US Blacks, the proportion older than 40 years was substantially greater (67 vs 48 per cent) and the distribution of educational attainment more extreme (31 per cent of policyholders versus 41 per cent of US Blacks had completed high school but not post-secondary education).*** However, the sociodemographic correlates of smoking in this survey (i.e., lower educational level, unemployment, being male, and being unmarried, poorer health status) are similar to those previously identified both for Black Americans[2,3,8] and for the US population as a whole.[2–4] It appears that the higher smoking prevalence among Blacks can be partially accounted for by Black/ White differences in these and other smoking-related sociodemographic factors.[3]

Present findings of a modal low-rate/high nicotine menthol smoking pattern also agree closely with existing data for US Black smokers. Black smokers, both men and women, smoke fewer cigarettes per day than White smokers;[2,4,5,17] in the present study, an estimated 52 per cent of smokers smoked 10 or fewer cigarettes per day. Comparable 1985 HIS figures were approximately 50 per cent for Black men and women smokers, a figure about twice as high as for non-Hispanic White smokers.[2] The 1981 Health Interview Survey found that 65 per cent of Black smokers smoked brands with 1.1 mg or more nicotine (in contrast to only 35 per cent of White smokers), and that 67 per cent of Black smokers smoked menthol cigarettes (in contrast to only 26 per cent of White smokers).[7] The present survey indicated very similar patterns: 63 per cent of smokers smoked brands with 1.0 mg or more nicotine; 73 per cent smoked menthol brands. Likewise, a recent regional study of 457 Black smokers found that three menthol brands (Kool, Newport, and Salem) accounted for nearly 60 per cent of all brands reported.[15] In the present survey, these same three brands accounted for 55 per cent of the brands reported.

Present evidence for relatively high nicotine dependency among Black smokers suggests that low-rate smokers of high nicotine menthol cigarettes may achieve higher than assumed levels of nicotine intake. Menthol cigarettes tend to be higher than non-menthol in tar and nicotine content.[15]

Sociodemographic correlates of quitting motivation in this Black survey population proved similar to those among US smokers in general, including: beliefs in smoking health harms and quitting benefits; recent smoking-related illness; personal medical advice to quit smoking; past efforts to quit or cut down; and expected social support for quitting.[18–20] Likewise, the quitting motives reported by ex-smokers are similar to those for national samples, with health concerns and self-mastery motivation foremost.[18–20] Hence, we would expect Black smokers to benefit from health education messages which emphasize the health harms/benefits of smoking/quitting, note that Blacks are not immune to these effects,[8] and are conveyed using sources, channels and media demonstrated to reach Black audiences.[7]

Past survey data suggest that Black smokers may try to quit as often as White smokers, but may succeed less often.[4,21] Present findings show relatively high quitting motivation and awareness of smoking health harms, and fairly active use of cut-down and self-quitting strategies among the current smokers surveyed.

On the other hand, present results show relatively limited use of beneficial quitting programs and resources by current or ex-smokers. There are, to date, no published data on the extent to which Black and White US smokers may differ in utilization of quit smoking services and resources. However, since Black Americans have more limited access to health care, particularly preventive care,[6] their access to quitting services and resources may also be more limited. Smoking cessation treatment outcomes for Black and White smokers of the same general socioeconomic levels are similar.[22,23] Therefore, improving access to existing treatments through established organizations within the Black community (e.g., churches, fraternal organizations, employers) should prove helpful. However, at the same time, it may be helpful in some circumstances to tailor the content and format of Black-focused treatments for maximum appeal and efficacy. Present findings suggest that such tailoring might:

- challenge possible misconceptions about the safety of low-rate smoking[24];
- introduce nicotine fading strategies designed to help smokers of high nicotine brands prepare for quitting[25];
- include Black ex-smokers as "peer coping models" to help enhance quitting motivation and efficacy[26];
- emphasize practical suggestions for avoiding excessive weight gain after quitting; and
- incorporate only print materials which have been pretested for comprehension and readability if the target audience includes a high proportion of smokers with limited formal education.[27]

***Unpublished tabulations, 1985 National Health Interview Survey (HIS), provided by the Office on Smoking and Health, US Public Health Service, Rockville, Maryland.

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American Lung Association Report Says EPA and the States Fail Clean Air Test

A new report from the American Lung Association (ALA) says the US Environmental Protection Agency (EPA) has failed to fulfill its mandate to protect the air Americans breathe. According to the report, EPA's "inadequate enforcement of the Clean Air Act" has allowed many states to weaken, delay, or fail to implement important air pollution control measures. Data were analyzed from nine metropolitan areas considered representative of major urban areas with a variety of air pollution problems from an array of sources.

The ALA report found "significant discrepancies" between many states' projections and actual achievements for air pollution reductions, citing as possible reasons: inaccurate computer models; overly optimistic projections of pollution reductions that would result from specific strategies; and deliberate noncompliance with clean air plans. ALA says that EPA has been lax in enforcing state air pollution control commitments, and states took advantage of it. In general, the states complied with the easier, less expensive controls, ALA pointed out, but not the tough measures Congress mandated.

The ALA report recommends that the next Congress authorize a new, stronger version of the Clean Air Act, giving both EPA and the states stricter guidelines and more resources. The report, *Compliance of Selected Areas with Clean Air Act Requirements*, also recommended that new or improved control measures be added to the Clean Air Act, i.e., stronger tailpipe emission standards; tighter national requirements for vehicle inspection and maintenance; and national regulations for area-wide sources such as dry cleaners, gas stations, architectural coatings, and aerosol sprays. For a full copy or executive summary of the report and a chart of the nine cities, contact American Lung Association, 1740 Broadway, New York, NY 10019–4374. Tel: (212) 315–8700.

A Randomized Trial of a Serialized Self-Help Smoking Cessation Program for Pregnant Women in an HMO

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Abstract: We report the results of a population-based randomized clinical trial that tested the effectiveness of a prenatal self-help smoking cessation program. The intervention consisted predominately of printed materials received through the mail. The population (n=242) consisted of a socioeconomically and ethnically diverse group of pregnant women enrolled in a large health maintenance organization (HMO) who reported they were smoking at the time of their first prenatal visit. Biochemical confirmation of continuous abstinence achieved prior to the 20th completed week of pregnancy and lasting through delivery revealed 22.2 per cent of the women in the eight-week serialized program quit versus 8.6 per cent of controls with usual care. The adjusted odds ratio was 2.80 (95 per cent CI=1.17, 6.69). We conclude that a low-cost prenatal self-help intervention can significantly affect the public health problem of smoking during pregnancy and its associated risks for maternal and child health. (*Am J Public Health* 1989; 79:182–187.)

Introduction

Smoking during pregnancy is an important modifiable risk factor for low birthweight,[1,2] and is also correlated with increased risk of early fetal loss.[3,4] It is projected that if all pregnant women stopped smoking, the number of fetal and infant deaths would be reduced by 10 per cent.[5] An estimate of the average cost of neonatal care is $189 higher (1983 dollars) for infants born to smokers.[6] Smoking in pregnancy has also been linked to postnatal complications.[7,8,9]

Some 25–30 per cent of US women smoke throughout their pregnancies.[10,11] The earlier a pregnant woman gives up smoking, the lower her excess risk of having a low birthweight infant; smoking cessation early in pregnancy reduces this risk to the same level as that of a non-smoker.[12] Studies conducted to date support the conclusion that higher prenatal quit rates can be achieved through intervention, although the majority of these findings have been limited by either weak methodology, small sample size, or reliance on self-reported outcomes.[13] Furthermore, most of these studies, including the two strongest,[14,15] have used an eighth month point-prevalence measure and therefore they have included some women who may have quit smoking too late in the pregnancy to achieve full benefit from cessation.[16]

This report presents the outcomes of a self-help smoking cessation program in a population receiving prenatal care through a health maintenance organization (HMO). The outcome measure is biochemically confirmed continuous abstinence beginning prior to the 20th week of pregnancy and lasting through delivery.

Methods

Sample and Design

The study design was a prospective randomized control group trial. Intake was conducted from July 1985 through June 1987 among English-speaking pregnant women who obtained prenatal care in one of five health centers of a large multispecialty group affiliated with Maxicare Health Plans, Inc., of Los Angeles, California.

All patients of the medical group began their prenatal care with an individual 45-minute conference conducted by a health educator. Of the total population of pregnant women seeking care, 95 per cent (n=2,382) met the study criterion of being less than 18 weeks pregnant at intake.

Smoking status was obtained at the start of the conference via an item on a self-administered screening questionnaire: "Which of the following statements best describes your smoking habits?" Response categories included: "I smoke now," "I smoke now, but I have cut down since I became pregnant," "I smoke from time to time," "I quit smoking since I became pregnant," and "I don't smoke." Of the 541 prepregnancy smokers (categories 1–4), the study was restricted to the 323 women still smoking ≥7 cigarettes/week. In a random sample of 132 reported non-smokers (category 5), only 4 per cent had biochemical values inconsistent with a non-smoking status.

At the conclusion of the conference, the health educator conducted a two-minute smoking-related interview with the smokers. All patients had been randomized in advance of their visit. Of the smokers, 165 were experimental (E) and 158 were control (C); the health educator was blind to group assignment until the end of data collection. Analyses for this paper are based on the 126 experimental and 116 control women who remained eligible for the study. Attrition resulted from abortion (E=7,C=11), miscarriage (E=12,C=13), disenrollment from the HMO or transfer to another medical group (E=20,C=18).

Intervention

After the smoking-related interview, in keeping with the standard practice of the medical group, all subjects were given a two-page pamphlet on the hazards of smoking during pregnancy and the importance of quitting. The health educator reinforced the written information in a two-minute discussion and answered any questions; she also advised patients of a five-session smoking cessation class available free through the HMO.

Experimental subjects were then given the first of eight self-help booklets, together with a three-minute overview of the program. Women were asked to make a commitment to read Booklet 1 within the ensuing week and to list their reasons for and against smoking. Each of the booklets had such activity assignments designed to personalize the program. The remaining seven booklets were mailed thereafter at weekly intervals and comprised a step-by-step program to

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increase motivation for quitting and to teach behavioral strategies for cessation and for relapse prevention. These four- to eight-page booklets, tailored to pregnancy, were printed in black ink on colored paper, had covers illustrated with cartoons depicting multi-ethnic figures, and were written at a ninth grade reading level.[17] They carried the logos of the HMO and medical group. All experimental group women received the program regardless of their motivation to quit smoking or desire to receive the program materials.

Patients had no further contact with the intake health educator. Prenatal care providers were blind to group assignment, and no effort was made to modify their usual counseling practices. Other sources of information on smoking—available through the HMO (e.g., childbirth classes)—were also left unchanged.

Measures

Data were from three sources: the intake conference, a telephone interview conducted at the 26th week of pregnancy, and urine samples routinely collected at each prenatal medical care visit.

The intake conference elicited a medical history, frequency of nausea, information about exposure to passive smoke, and beliefs about the effect of smoking on fetal and maternal health. Additional questions asked of smokers included a brief smoking history, smoking in the current pregnancy, smoking social networks and support, seriousness about quitting, likelihood of quitting, and degree of confidence in ability to quit (Table 1).

The 26th week telephone interview was completed with 94 per cent of the experimental group and 92 per cent of the controls. A standard interview schedule was used to collect data on health-related behavior during the pregnancy. The smoking section, embedded in questions about these other behaviors, included smoking in the past seven days, quit attempts, length of time off cigarettes, and number of slips for those not currently smoking. The interview also assessed the degree to which the experimental group read and used the booklets, satisfaction with content and appearance, and perceived usefulness of the self-help program.

Medical charts were flagged to prompt nurses to retain the urine samples of the study subjects. Patients were usually seen monthly until their ninth month when visits were scheduled each week. On average, seven urine samples were obtained for each subject. After delivery, selected samples were biochemically analyzed for the presence of cotinine, a derivative of nicotine.[18]

Quits reported in the telephone interview were confirmed with all relevant urine samples. A quit was defined as a minimum of seven consecutive days without smoking "even a puff." Additionally, all subjects had a sample from the 34th week of pregnancy analyzed to verify maintenance of a quit or to detect a quit which occurred after the telephone interview. Subjects were classified into the categories of Early Quitter, Middle Quitter, Late Quitter, Early Relapser, Late Relapser, and Non-Quitter. Of primary interest in the current analysis is the group of Early Quitters. These were women who sustained a quit which occurred prior to the 20th week of pregnancy and therefore could derive the greatest benefit from quitting smoking. On average, an Early Quit was based on three urine samples verifying a non-smoking status.

TABLE 1—Comparison of Experimental Conditions: Background Characteristics and Selected Covariates

	% Experimental (126)	% Control (116)	% Total (242)
Sociodemographics			
Race			
White	65.9	62.1	64.0
Black	25.4	26.7	26.0
Other	8.7	11.2	10.0
Education			
Less Than High School	16.7	19.8	18.2
High School-Some College	73.8	72.4	73.1
College Graduate	9.5	7.8	8.7
Marital Status			
Married	60.3	59.5	59.9
Age (years)			
18–19	7.2	9.5	8.3
20–29	71.4	59.5	65.7
30+	21.4	31.0	26.0
Pregnancy-Related			
Gravida			
0	20.6	17.2	19.0
1	27.8	18.1	23.1
2+	51.6	64.7	57.9
Number of Children			
0	50.0	41.4	45.9
1	33.3	30.2	31.8
2+	16.7	28.4	22.3
Weeks Entering Care			
<9	31.0	31.9	31.4
9–13	50.0	44.8	47.5
14+	19.0	23.3	21.1
Nausea Experienced			
Very Often/Often	46.8	47.4	47.1
Not Too Often	34.9	40.5	37.6
Not At All	18.3	12.1	15.3

(continued)

TABLE 1—Continued

	% Experimental (126)	% Control (116)	% Total (242)
Smoking History			
Age Began Smoking (years)			
Less Than 16	37.3	31.9	34.7
16–18	46.8	45.7	46.3
19+	15.9	22.4	19.0
Prepregnancy Smoking Rate			
1–10 cgts/day	30.2	24.1	27.3
11–19	13.5	14.7	14.0
20+	56.3	61.2	58.7
Minutes to First Cigarette			
1–20	27.0	36.3	31.4
21–60	35.7	28.4	32.2
61+	37.3	35.3	36.4
Previous Serious Quit Attempts			
0	22.2	29.3	25.6
1	30.2	29.3	29.8
2+	47.6	41.4	44.6
Longest Time off Cigarettes			
0 Days	13.5	16.4	14.9
1–7 Days	31.7	33.6	32.6
8+ Days	54.8	50.0	52.5
Pregnancy-Specific Smoking Behavior			
Smoking Rate at Intake			
1–10 cgts/day	72.2	71.6	71.9
11–19	13.5	16.4	14.9
20+	14.3	12.1	13.2
Cut Down from Prepregnancy Rate			
Yes	78.6	79.3	78.9
Quit for at Least a Day			
Yes	31.7	28.4	30.2
Pregnancy-Specific Smoking Health Beliefs and Attitudes			
Seriousness about Quitting			
Very	62.7	56.0	59.5
Somewhat	27.8	32.8	30.2
Not Too/Not At All	9.5	11.2	10.3
Likelihood of Quitting			
Very	52.4	44.8	48.8
Somewhat	29.4	33.6	31.4
Not Too/Not At All	18.2	21.6	19.8
Confidence in Ability to Quit			
Very	34.9	31.9	33.5
Somewhat	41.3	37.1	39.3
Not Too/Not At All	23.8	31.0	27.2
Smoking during Pregnancy Harms Baby			
A lot	44.4	49.1	46.7
Somewhat	37.3	32.8	35.1
Little/None/Don't Know	18.3	18.1	18.2
Smoking during Pregnancy Harms Mother			
A lot	37.3	43.1	40.1
Somewhat	41.3	34.5	38.0
Little/None/Don't Know	21.4	22.4	21.9
Smoking Social Networks/Social Support			
Smoking Status of Partner*			
Partner Smokes	52.4	67.2	59.5
Number of Other Smokers in Household			
1+	59.5	69.0	64.0
Smoking Status of Friends			
All/Most	37.3	40.5	38.8
Some	46.8	46.6	46.7
None	15.9	12.9	14.5
Anticipated Help in Effort to Quit			
A Great Deal	58.7	52.6	55.8
Somewhat	21.4	25.0	23.1
Little/None	19.9	22.4	21.1

*P<.05

TABLE 2—Smoking Cessation Outcome Classifications

Quit Category	Self-reported Smoking Status at Interview[a]	Selection of Urine Samples Analyzed	Cotinine Cutoffs[b]
Early Quitter	Not currently smoking and the quit began when ≤20 weeks pregnant[c]	All available samples from start of quit through interview + 34th week of pregnancy	At least one value ≤10 ng/ml and no value ≥30 ng/ml[d] + ≤30 ng/ml
Middle Quitter	Not currently smoking and the quit began when >20 and ≤26 weeks pregnant	All available samples from start of quit through interview + 34th week of pregnancy	At least one value ≤10 ng/ml and no value ≥30 ng/ml[d] + ≤30 ng/ml
Late Quitter	Currently smoking and no quit prior to interview	No sample prior to interview 34th week of pregnancy + 35th week of pregnancy[e]	≤10 ng/ml + ≤10 ng/ml
Early Relapser	Currently smoking and a quit prior to interview	All available urines during quit + 34th week of pregnancy	At least one value ≤10 ng/ml and no value ≥30 ng/ml[d] + ≥30 ng/ml
Late Relapser	Currently not smoking and a quit prior to the interview	All available samples from quit through interview + 34th week of pregnancy	At least one value ≤10 ng/ml and no value ≥30 ng/ml[d] + ≥30 ng/ml
Non-Quitter[f]	Currently smoking and no quit prior to interview	No sample prior to interview 34th week of pregnancy	>10 ng/ml

[a]Subjects were asked about their smoking status in a telephone interview conducted in the 26th week of their pregnancy. For the 7% of cases without an interview, classification was based on cotinine values from all available urine samples prior to the 26th week of pregnancy, as well as the 34th week.

[b]Smoking status was determined by assay of urine samples using a version of the gas-liquid chromatography method with nitrogen-phosphorous detection for cotinine.

[c]On average, Early Quitters began their sustained quits in the 11.9th week of pregnancy (S.D. 3.3).

[d]For subjects reporting no puffs during a quit and low exposure to passive smoke, all values prior to the interview had to be ≤10 ng/ml. For subjects reporting at least one slip during the quit or high exposure to passive smoke, some values could be as high as 29 ng/ml although at least one had to be ≤10 ng/ml. (Exposure to passive smoke was assessed by an item from the screening questionnaire: "How much are you exposed to the cigarette smoking of others?" Response categories ranged from 0 "No one smokes around me" to 100 "Others smoke around me all the time." High = 100 for combined responses for home and work.

[e]Due to the self-reported smoking status at interview, a subject was classified as a Late Quitter only if two consecutive urine samples had cotinine values ≤10 ng/ml.

[f]Self-reported quitters with cotinine values which exceeded cutoffs were included in this category (approximately 10% of self-reported quits).

(See Table 2 for a detailed description of the classification methodology.)

Results

Table 1 shows that with the exception of partner's smoking status, the experimental and control groups were similar in background characteristics and in variables with a hypothesized relationship to cessation. The subjects were predominately White, married, in their twenties, and had at least a high school education (mean 12.5 years). The majority had two or more previous pregnancies and entered care for the current pregnancy in the first trimester. At the time of the intake conference, about three-fourths of the sample were smoking ≤10 cigarettes/day with the mean daily consumption half the prepregnancy rate (18.7 versus 9.1). The majority, however, had not quit smoking for as long as 24 hours since becoming pregnant.

Most women were serious about wanting to quit during their current pregnancy but somewhat less confident in their ability to do so. Most women believed that smoking during pregnancy could harm either their own or their baby's health. The majority reported that their partner smoked and about half expected to receive "a great deal" of help if they decided to quit smoking.

Smoking Cessation Outcomes

Table 3 contrasts the study groups by biochemically confirmed category of quitting. The self-help program yielded more than two times the proportion of Early Quitters than

TABLE 3—Smoking Cessation Outcomes by Experimental Status

	% Experimental (N=126)	% Control (N=116)	Difference	95% CI
Early Quitter	22.2	8.6	11.5*	03.0,20.0
Middle Quitter	0.8	1.7		
Late Quitter	3.2	6.9		
Early Relapser	1.6	0.0		
Late Relapser	0.8	0.0		
Non-Quitter	71.4	82.8		

*Adjusted for group differences in prepregnancy smoking rate, likelihood of quitting, number of children, belief of smoking's harm to the baby, number of smokers in household, and smoking status of partner.

TABLE 4—Early Quit Rates by Experimental Condition (by selected covariates)

Total Sample	% Experimental (22.2)	% Control (8.6)	% Total (15.7)
Prepregnancy Smoking Rate			
1–10 cgts/day	42.1	21.4	33.3
11–19	35.3	5.9	20.6
20+	8.5	4.2	6.3
Week Entering Care			
<9	17.5	13.5	15.6
9–13	27.0	7.7	18.3
14+	17.4	3.7	10.0
Number of Children			
0	25.8	14.9	21.1
1	20.9	8.6	15.4
2	14.3	0.0	5.5
Confidence in Ability to Quit			
A Great Deal	31.8	10.8	22.2
Somewhat	23.1	11.6	17.9
Little/None	6.7	2.8	4.5
Number of Smokers in Household			
0	25.5	13.9	20.7
1+	20.0	6.3	12.9

was observed among controls. Adjusting for the covariates specified in Table 3, the odds ratio was 2.80 (95 per cent CI = 1.17, 6.69).

Late Quitters (subjects who quit in the third trimester) comprised the majority of the other biochemically confirmed quits, providing evidence that some sustained quits do begin close to the end of pregnancy. Only 1 per cent of the total sample was classified as relapsers.

We next examined group proportions of Early Quitters within the context of selected covariates to determine whether the intervention had differential impact for particular subgroups of pregnant women (Table 4). The percentage difference between experimental conditions was about five times as great among light smokers (subjects with a prepregnancy smoking rate <20 cigarettes/day) compared to the heavier smokers. A similar five-fold difference was observed within strata of confidence in ability to quit. In a log-linear model,[19] none of the five interactions was statistically significant at the $p = .05$ level. The failure to detect subgroup differences, however, must be viewed within the context of the very limited power to detect even moderately large interactions.

Process Evaluation

Table 5 provides data on the experimental group's use and evaluation of the program. The vast majority reported reading at least one of the eight booklets, with a mean of 5.8 booklets read, and nearly half (47.7 per cent) reported reading the entire series. The booklets were judged to be interesting, attractive, helpful, and easy to understand. Half of the Early Quitters indicated that the booklets influenced their decision to quit, although a third rated the booklets of "no help at all."

TABLE 5—Use and Evaluation of Self-Help Smoking Program

Use/Evaluation	Mean	SD
Read At Least One Booklet (93%)		
Number of Booklets Read: (1–8)	5.78	2.36
How Thoroughly Read*	6.85	2.58
How Easy They Were to Understand*	9.42	1.14
How Interesting They Were*	7.97	2.06
Their Attractiveness*	7.68	1.98
Their Helpfulness*	7.98	2.64
How New the Information Was*	7.07	2.83

*Scale of 1–10
94% of Experimental Group Responding.

Discussion

Findings from this study indicate that an eight-part series of smoking cessation booklets addressed to the issues of pregnancy were effective in assisting an ethnically and socioeconomically diverse group of pregnant women to quit smoking.

The program raised Early Quit rates in all of the subgroups examined, including those in which women were not likely to quit smoking on their own (those entering care in the second trimester, multiparous women, and those with adverse smoking social networks). While the sample size precluded precise estimation of study effects within subgroups, the observation that the intervention was most effective among women with light prepregnancy smoking rates or with initial high levels of self-efficacy is plausible.

Although our primary interest was in continuous abstinence, other relevant studies to date have measured outcomes with an eighth-month point prevalence. Our study design precluded duplication of this measure; however, combining the Early, Middle and Late Quitter categories provides a close approximation. Using these criteria we found a quit rate of 26.2 per cent in the experimental group, similar to the results achieved in two other trials using rigorous study designs among similar populations. Sexton achieved a quit rate of 32 per cent with an intervention that included one personal visit with a health educator at the time of enrollment, a monthly phone call, and biweekly contact by mail usually in the form of a newsletter.[20] Secker-Walker, *et al*, achieved a cessation rate of 24 per cent among subjects exposed to three individual, 10-to-15-minute counseling sessions with a nurse or health educator; subjects were also given informational materials.*

While it is difficult to draw exact comparisons between studies due to differing patient characteristics, recruitment

*Flynn B: Personal communication, May 1988.

strategies, and practice settings, it appears that this program which used self-help achieved results comparable to those which relied more heavily on personal contact. In contrast to counseling-based programs, self-help interventions in general are easier to implement, can reach a larger target population, and are less costly.

The content of our booklets was similar to other self-help programs; however, the nature of its delivery was unique. Our clinical impression is that the weekly mailings helped keep the women actively involved in their attempts to quit smoking through regular cues to action. Furthermore, women who initially may have been unreceptive to an intervention were afforded an opportunity to benefit from the booklets if their motivation to quit changed during the two-month period.

Another issue not empirically examined but which we believe was helpful was the personalized introduction of the program as a routine part of prenatal care. Women were prescribed the smoking cessation program just as they were their prenatal vitamins, thereby heightening the credibility and impact of the materials.

We have estimated the total costs of the intervention to be approximately $11/patient as implemented in this study (i.e., development and printing costs for 2500 booklets, personnel time and mailing). Had the entire program been provided at one time without a personal introduction, the cost would have been $4/patient.

To date, prenatal smoking cessation interventions in non-welfare populations have yielded quit rates in the range of 25 per cent. Of note, approximately half of the Non-Quitters reported at the 26th week interview that they had quit for at least a day. Therefore, a promising avenue for future research is the development of interventions to convert these unsuccessful attempts into sustained quits.

The model of group-based smoking cessation classes has not proved successful in reaching a wide target population of pregnant smokers. Hughes, *et al*,[21] working with patients receiving care in a hospital outpatient clinic, and Windsor[15] with patients in a public maternity clinic, were unable to induce women to attend such classes. Similarly, in our HMO population we found only 2.4 per cent of patients attended the free classes. In contrast, a self-help program can reach all pregnant smokers.

The simplicity of our program permits its implementation in a variety of health care settings such as medical groups, individual practices, and maternity clinics. The self-help design allows its introduction by a range of medical staff personnel from paraprofessionals to obstetricians. This program offers another effective method of addressing the major public health problem of smoking during pregnancy. The challenge remains for the widespread incorporation of smoking cessation interventions into standard prenatal care.

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Compensation as a Risk Factor for Lung Cancer in Smokers Who Switch from Nonfilter to Filter Cigarettes

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Abstract: The likelihood of increasing the number of cigarettes per day (cpd), a common method of compensation, in smokers switching from nonfilter to filter cigarettes, was examined in newly diagnosed and histologically confirmed lung cancer cases including all cell types: 1,242 lung cancer cases and 2,300 sex and age matched hospital controls were interviewed in 20 hospitals from nine United States cities between 1969 and 1984.

The mean increase in cpd for lung cancer cases was about twice that of controls. Using switchers who did not increase cpd as the referent group, there was a linear dose-response relationship between the odds of lung cancer and increasing level of compensation. The odds ratio increased from 1.19 to 2.37 in males and from 1.66 to 3.83 in females corresponding to increases of 1–10 to 21+ cpd after switching.

Findings from this study suggest that increasing cpd after switching to filter cigarettes is an important risk factor for lung cancer that needs to be emphasized in epidemiologic studies. Proponents of the idea that switching cigarettes is of some benefit, should also advocate that individuals who continue smoking should avoid compensation after switching. (*Am J Public Health* 1989; 79:188–191.)

Introduction

Epidemiologic studies have shown a reduced risk for lung cancer among smokers who switch from nonfilter to filter cigarettes relative to those who continue smoking nonfilter cigarettes.[1] On the basis of this finding it has often been recommended that people who are unwilling or unable to give up smoking should switch to filter cigarettes. The tar intake per day after switching to filter cigarettes will be reduced if smokers continue to smoke the same or fewer numbers of cigarettes per day (cpd). In this paper the term "switchers" stands explicitly for smokers who switched from nonfilter to filter cigarettes. The ways by which switchers attempt to compensate for the reduction in nicotine are by increasing cpd or altering their puff frequency, puff duration, puff rate, and interpuff interval[2] exposing themselves to higher quantities of gas and particulate cigarette smoke constituents.[3] This study focuses on compensation in switchers who increase cpd among lung cancer cases and controls. The main questions of interest were: Is there a differential pattern of increase in cpd between cases and controls? Is there an increase in risk for lung cancer in switchers who increase cpd compared to those who do not?

Methods

The cases and controls in this study were interviewed between 1969 and 1984, in 20 hospitals in nine US cities, as part of a hospital-based ongoing study of smoking-related cancers.[4] In the original study, cases were patients with newly diagnosed and histologically confirmed tobacco-related cancers and controls were hospitalized patients with an admitting diagnosis consisting of conditions not thought to be related to smoking. The exclusion of smoking-related conditions from control diagnoses was intended to reduce potential bias in the assessment of risks due to cigarette smoking. For the purpose of this study, we took lung cancer cases of all cell types and controls from this data pool. Controls had cancers of other organs and non-neoplastic diseases. The smoking-related cancers excluded from control diagnoses were cancers of the oral cavity, larynx, nasopharynx, kidney, bladder, oesophagus, salivary gland, pancreas and liver; patients with heart disease, stroke, peripheral vascular disease, chronic bronchitis, emphysema, gastrointestinal ulcer, and cirrhosis were also excluded.

All study subjects were interviewed in the hospital by trained interviewers using a structured questionnaire that included sociodemographic variables, lifetime smoking history, cpd, brands of filter and nonfilter cigarettes smoked. For each individual, the average cpd of filter cigarettes was calculated by summing up the cpd for all filter brands and then dividing by the number of filter brands. The average cpd of nonfilter cigarettes was calculated similarly.

As a partial validation procedure, we checked the reliability of the initial smoking data by recontacting 100 cases and 100 controls, selected at random following their discharge from hospital.

In this investigation, only those who were current cigarette smokers and who had switched from nonfilter to filter cigarettes were considered. A current smoker was defined as someone who had smoked at least one cpd for a continuous period of one year or more and was also smoking within the year preceding diagnosis. Those cigarette smokers who smoked cigars or pipes concurrently were excluded. Cases were frequency matched at approximately a 1:2 ratio on sex and age at diagnosis (± 5 years) to controls.[5]

Statistical Techniques

All statistical analyses were performed separately by sex. Paired cpd data before and after switching were used to assess the change in cpd. Mean changes in cpd after switching for cases and controls were adjusted by linear regression for age at switching and duration of nonfilter cigarette smoking utilizing analysis of covariance.[6] Adjusted means for cases and controls were compared by t-tests using the error mean square from the analysis of covariance with the appropriate degrees of freedom.[6,7]

Logistic regression analysis[8] was used to estimate the odds ratio for lung cancer in those who compensated (increased cpd) relative to those who did not compensate. An unmatched analysis was performed with adjustments for nonfilter cpd, duration of nonfilter and filter cigarette smoking, and age at diagnosis or age at switching. Since our study was limited to switchers, all subjects had corresponding sets of time points in their smoking histories (age at start, age at

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switch, and age at diagnosis), thereby facilitating simultaneous adjustment for the same potential confounders in multiple logistic models. In this paper, we restrict the term "compensation" to mean an increase in cpd after switching to filter cigarettes.

Multiple linear regression analysis[6] was performed among controls to determine if cpd after switching could be predicted by variables such as age at switching, nonfilter cpd, nonfilter duration, and tar yield per day before switching.

Results

A total of 781 male and 461 female lung cancer cases of all cell types and 1,432 male and 868 female controls were studied; all subjects were smokers who had switched to filter cigarettes. Among male switchers, 45 per cent (n = 351) of cases and 41 per cent (n = 580) of controls increased cpd; 42 per cent (n = 327) of cases and 44 per cent (n = 634) of controls smoked the same number of cpd; and 13 per cent (n = 103) of cases and 15 per cent (n = 218) of controls decreased cpd. Among female switchers, 59 per cent of cases (n = 272) and 48 per cent of controls (n = 418) increased cpd; 31 per cent (n = 143) of cases and 41 per cent (n = 356) of controls smoked the same number of cpd; and 10 per cent (n = 46) of cases and 11 per cent (n = 91) of controls decreased cpd.

The maximum reduction after switching among cases and controls was about 50 cpd in males and 60 cpd in females. The maximum increase noted ranged from 50–70 cpd in males and 50–60 cpd in females. In both males and females, the interquartiles (25th to 75th percentiles) ranged from 0–15 cpd in cases and 0–10 in controls.

Distributions of cases and controls by age, education, marital status, religion, and occupation are presented in Table 1. In both males and females, case-control differences were not evident in the sociodemographic variables. Mean ages at diagnoses for males were 57.0 years in cases and 56.1 years in controls. The corresponding ages for females were 56.8 years and 55.8 years.

Mean values of duration of smoking and cpd are shown in Table 2. On the average, both male and female cases had smoked nonfilter cigarettes 1.5–2 years longer than controls. In males, cases and controls smoked approximately the same duration of filter cigarettes while female cases smoked approximately one year longer than the controls. On the average, cases of both sexes smoked about 5 to 6 more cpd than the controls before switching, and about 8 more cpd than the controls after switching. Notably cases showed greater compensation than controls, the mean increase in cpd for cases being about twice that of controls. Adjustments for the differences between cases and controls in duration of nonfilter cigarette smoking, age at switching and age at diagnosis by analysis of covariance did not alter these results.

Mean values of cpd before and after switching by cpd classes before switching (1–10, 11–20, 21–30, 31–40 and 41+) are presented in Table 3. Three points are evident from these data. Most importantly, cases showed greater average compensation than the controls in each cpd class and the differences in compensation between cases and controls within classes were especially great for all the classes in males and for the first four classes in females. Compensation in cpd was evident only when the initial cpd was less than 30. When initial cpd was greater than 30, there was a tendency either not to increase or even to decrease cpd after switching. In both cases and controls, the extent of compensation progressively declined with increasing cpd before switching. The difference in cpd after adjustments using analysis of covariance for age at switching and duration of nonfilter cigarette smoking did not alter these results appreciably.

Table 4 shows a clearcut dose-response relationship for lung cancer with increasing cpd levels (compensation) in both

TABLE 1—Distribution of Demographic Variables among Cases and Controls

Demographic Variables	Males		Females	
	Cases N = 781 %	Controls N = 1432 %	Cases N = 461 %	Controls N = 868 %
Age (years)				
<44	7.8	8.8	10.2	10.6
45–54	29.1	33.2	27.8	31.5
55–64	42.1	41.5	41.2	40.9
65–74	19.6	15.2	18.2	15.7
75+	1.4	1.3	2.6	1.4
Education				
None, Grammar School	21.0	19.9	12.8	13.2
<High School	24.5	24.0	25.0	21.3
High School Graduate	28.3	26.4	38.0	37.5
Some College	11.3	15.9	15.4	18.0
College	9.6	8.5	5.4	6.7
Postgraduate	5.4	5.3	3.5	3.3
Marital Status				
Single	6.0	7.7	6.7	6.3
Married	74.1	70.0	53.4	55.9
Divorced	9.5	13.0	12.2	12.0
Separated	4.4	4.3	4.3	5.7
Widowed	5.9	5.0	23.4	20.1
Religious Status				
Protestant	42.3	49.7	33.7	29.3
Catholic	43.8	34.5	44.1	36.7
Jewish	9.1	9.5	17.2	13.9
Other	2.2	3.2	3.0	3.4
None	2.6	3.2	2.0	1.2
Race				
Caucasian	85.2	81.8	91.5	85.6
Black	13.4	15.9	7.8	13.0
Hispanic	1.2	1.7	0.7	1.0
Oriental	0.3	0.4	0.0	0.2
Other	—	0.2	—	0.1
Occupation				
Professional	8.6	9.1	4.6	6.4
Business Executive	10.7	9.8	3.9	3.4
Technical	9.3	6.7	26.4	15.2
Clerical-Sales	21.2	20.4	8.3	12.7
Skilled	18.3	17.3	9.4	8.7
Semi-skilled	12.9	11.6	3.1	4.1
Unskilled	6.7	9.0	3.1	4.5
Retired Unemployed	17.0	16.1	6.5	11.1
Housewife	—	—	34.9	34.0

TABLE 2—Mean Duration of Smoking and Cigarettes per Day (CPD) Smoked Before and After Switching among Cases and Controls

Total Number	Males		Females	
	Cases N = 781	Controls N = 1432	Cases N = 461	Controls N = 868
Mean Nonfilter duration in years (SE)	26.3 (0.4)	24.8 (0.3)	22.1 (0.6)	19.8 (0.4)
Mean Filter duration in years (SE)	13.5 (0.3)	13.0 (0.2)	15.6 (0.4)	14.3 (0.3)
Mean Nonfilter cpd (SE)	28.5 (0.5)	22.7 (0.3)	21.8 (0.5)	16.8 (0.3)
Mean Filter cpd (SE)	34.4 (0.6)	26.6 (0.4)	29.7 (0.6)	21.5 (0.4)
Difference in cpd[a] (SE)	5.9 (0.5)	3.9 (0.3)	7.8 (0.6)	4.7 (0.3)

[a]Difference in cpd = Filter cpd − Nonfilter cpd.
SE = Standard Error

TABLE 3—Mean Cigarettes per Day (CPD) Smoked Before and After Switching According to Nonfilter CPD Groups

	Nonfilter cpd in Males									
	1–10		11–20		21–30		31–40		41+	
	Cases n = 42	Controls n = 212	Cases n = 324	Controls n = 786	Cases n = 174	Controls n = 208	Cases n = 158	Controls n = 163	Cases n = 74	Controls n = 57
Mean Nonfilter cpd (SE)	8.4 (0.4)	7.9 (0.2)	19.3 (0.1)	19.1 (0.1)	28.4 (0.2)	28.5 (0.2)	39.2 (0.2)	39.5 (0.1)	57.2 (1.0)	57.5 (1.3)
Mean Filter cpd (SE)	19.1 (1.9)	14.7 (0.7)	27.6 (0.7)	24.2 (0.4)	35.2 (1.0)	33.1 (0.8)	42.6 (1.1)	38.4 (1.2)	53.5 (2.1)	46.7 (2.5)
Difference in cpd[a] (SE)	10.7 (1.9)	6.8 (0.7)	8.2 (0.7)	5.0 (0.4)	6.7 (1.0)	4.6 (0.8)	3.3 (1.1)	−1.1 (1.2)	−3.7 (1.9)	−10.8 (2.3)
	Nonfilter cpd in Females									
	n = 87	n = 308	n = 241	n = 446	n = 72	n = 60	n = 39	n = 35	n = 20	n = 13
Mean Nonfilter cpd (SE)	8.4 (0.3)	7.6 (0.2)	19.0 (0.2)	18.6 (0.1)	28.0 (0.3)	28.3 (0.3)	39.1 (0.3)	39.0 (0.4)	56.9 (2.4)	58.5 (2.1)
Mean Filter cpd (SE)	22.4 (1.2)	14.5 (0.5)	27.1 (0.7)	22.9 (0.4)	35.3 (1.2)	31.7 (1.3)	42.5 (1.6)	36.6 (2.2)	45.7 (3.5)	49.7 (5.6)
Difference in cpd[a] (SE)	14.1 (1.3)	6.9 (0.5)	8.0 (0.7)	4.2 (0.4)	7.3 (1.2)	3.3 (1.3)	3.4 (1.6)	−2.4 (2.2)	−11.2 (4.5)	−8.8 (6.6)

[a]Difference in cpd = Filter cpd − Nonfilter cpd

sexes. The odds ratios increased from 1.19 to 2.37 in males and 1.66 to 3.83 in females corresponding with increases of 1–10 to 21+ cpd after switching.

The odds ratio associated with increasing cpd versus decreasing cpd after switching, relative to those who did not change cpd, was also examined. Compared to those who did not change their cpd, those who increased cpd had an odds ratio of 1.38 (95 per cent confidence intervals = 1.31, 1.69) in males and 2.19 (95 per cent CI = 1.67, 2.89) in females after adjusting for confounders. In contrast, males who decreased cpd had a reduction in the odds ratio to 0.64 (95 per cent CI = 0.47, 0.86); in females, the number of cases who decreased cpd was small (n = 46) and the corresponding odds ratio was 0.97 (95 per cent CI = 0.62, 1.55).

Since greater compensation occurred among those who initially smoked ≤30 cpd (Table 3), we examined the risks associated with increasing cpd according to those who initially smoked ≤30 cpd and >30 cpd. Among those who initially smoked ≤30 cpd, those who increased cpd had elevated odds ratios of 1.34 (95 per cent CI = 1.08, 1.65) in males, and 1.84 (95 per cent CI = 1.42, 2.38) in females for lung cancer, relative to those who did not increase cpd, after adjustments for age at switch, duration of nonfilter and filter cigarette smoking. Among those who initially smoked >30 cpd, the corresponding odds ratio was also elevated in males: OR = 1.67 (95 per cent CI = 1.08, 2.57); in females, OR = 0.94 (95 per cent CI = 0.37, 2.41); the numbers of cases (n = 59) and controls (n = 48) were relatively small.

TABLE 4—Logistic Regression Analysis Results Showing Risks of Lung Cancer for Those Who Increased CPD after Switching

Increase in Filter CPD	Cases	Controls	Odds Ratio[ab]	95% Confidence Interval
Males				
Did not increase cpd*	430	852	1.00	—
1–10 vs did not increase cpd	150	326	1.19	0.93–1.51
11–20 vs did not increase cpd	128	180	1.75	1.33–2.29
21+ vs did not increase cpd	73	73	2.37	1.64–3.41
Females				
Did not increase cpd*	189	447	1.00	—
1–10 vs did not increase cpd	126	262	1.66	1.23–2.24
11–20 vs did not increase cpd	101	115	2.97	2.09–4.20
21+ vs did not increase cpd	45	40	3.83	2.31–6.34

[a]Adjusted for nonfilter duration, filter duration, nonfilter cpd, and age at switch.
[b]Test of linear trend significant at $p < 0.05$.
*Referent category.

The regression of compensation (filter cpd − nonfilter cpd) on nonfilter cpd yielded negative slopes, indicating that nonfilter cpd was a negative predictor of compensation, i.e., the lower the initial cpd, the greater the compensation.

Linear regression analysis among controls was performed to examine the effect of selected variables in predicting cpd after switching. Results were similar for males and females. Cigarettes per day before switching ($\beta = 0.7$, S.E. = 0.03) and age at switching ($\beta = 0.2$, S.E. = 0.03) predicted cpd after switching in linear regression models. These variables together accounted for 34 per cent and 39 per cent of the variability among males and females, respectively, while duration of nonfilter cigarette smoking, tar yield per day before switching, and age at start of smoking accounted for little or none of the variability.

Discussion

The results of this study indicate that a sizable portion of smokers tend to compensate (increase cpd) after switching from nonfilter to filter cigarettes and that patients with lung cancer compensate more than controls. A greater proportion of females (50–60 per cent) than males (40–45 per cent) showed compensation in cpd. Among switchers who increased cpd, there was a linear dose-response relation between the odds of lung cancer with increasing level of compensation. The linearity of the dose response pattern in both sexes supports the biological plausibility of compensation as an important risk factor for lung cancer. We emphasize that these results are adjusted for past smoking patterns and duration differences.

In our study, the lower the initial cpd, the greater the tendency to compensate by increasing cpd. In fact, heavy smokers of nonfilter cigarettes (>30 cpd) on the average tended to decrease rather than increase cpd after switching to filter cigarettes. There may be several explanations for this finding. It is possible that heavy smokers may compensate by altering their smoking topography since there may be a time limitation (less leeway) in smoking additional cigarettes and/or encumbent symptoms might compel them to smoke less upon switching. A similar finding was noted in the 13-year follow-up survey by the American Cancer Society.[9] Lighter smokers (not separated by switch status) of less than half a pack a day in 1959 showed an average increase in cpd by 1972 which doubled that observed in heavier smokers. Another plausible explanation relates to an upper threshold phenomenon of nicotine dependency. In the already heavy nonfilter cigarette smokers (>30 cpd), the amount of filter cigarettes smoked (even when it is less than the initial cpd) after switching may be sufficient to maintain their required nicotine level.

One possible source of bias in this study is recall bias. If cases overestimated their exposure, while controls underestimated it, then the risk will appear larger than real. As a partial validation procedure, we reinterviewed 100 cases and 100 controls following their discharge from hospital. The results from the second interview showed good agreement with first reported cpd in both groups (average correlation 0.90) and revealed no differential reporting between cases and controls. The average difference between first and second interviews for both cases and controls was only one cpd.

The major reason for limiting our analysis to switchers only was to facilitate adjustment for the corresponding sets of confounders (duration of nonfilter smoking, age at switch, and duration of filter smoking) within the comparison groups, i.e., switchers who did not compensate, using multiple logistic analysis. Since there is also intense interest in comparing switchers with exclusive nonfilter cigarette smokers, we attempted to assess their patterns of odds ratio. Wynder and Kabat[10], observed minimum odds ratios for lung cancer of 0.66 and 0.74 in males and females, respectively, among switchers who had smoked filter cigarettes for more than 10 years, relative to exclusive nonfilter smokers. Assuming that 0.66 and 0.74 represent the odds ratios among male and female switchers who increased by 1–10 cpd after switching (this was the median class of cpd change in our study), we computed odds ratios due to various levels of compensation relative to nonfilter cigarette smokers (Figure 1). Further reductions in odds ratios are evident with no compensation (0.55 and 0.45 in males and females, respectively), whereas for switchers who compensate, the odds ratios increase in a linear fashion, and even become greater than that of nonfilter only smokers when the compensation level exceeds 20 cpd.

Compensation as a risk factor has not been stressed in epidemiologic studies comparing the health risk of nonfilter to filter cigarette switchers relative to exclusive nonfilter cigarette smokers. Compensation is undoubtedly a complex phenomenon. It includes changes in puff frequency, duration, rate and interval[2]. Nevertheless, the relatively simple measure of compensation used in this study (increase in cpd after switching) has proved to be an important risk factor for lung cancer in itself. Based on our results, the lung cancer risk among smokers who switch to filter cigarettes would be reduced further if they strive not to increase cpd after switching. Proponents of the idea that switching cigarettes is of some benefit should also advocate that individuals who continue smoking should avoid compensating after switching.

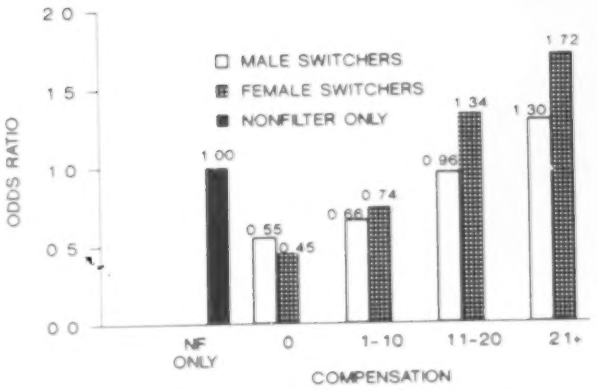

FIGURE 1—Risks among Switchers Relative to Nonfilter Only Smokers According to Different Levels of Compensation.*

*Compensation is the increase in cigarettes per day after switching to filter cigarettes.

SOURCE: Computed from the results of this study and the results of the study by Wynder and Kabat[10].

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A Comparative Evaluation of a Restrictive Smoking Policy in a General Hospital

Lois Biener, PhD, David B. Abrams, PhD, Michael J. Follick, PhD, and Larry Dean, BA

Abstract: The impact of a restrictive smoking policy on the behavior and attitudes of smokers and non-smokers was assessed by surveying random cross-sectional samples of hospital employees before, six months after and 12 months after the policy was implemented, and comparing responses with those of employees of a hospital with no restrictive policy. Effectiveness of policy implementation was also evaluated. Results indicated that the policy was well-publicized and was approved by virtually all the non-smokers and the majority of the smokers. Following implementation, employees in the smoking policy hospital were less likely to report being bothered by smoke at their work stations than were employees of the comparison hospital. Six months and one year after the policy change, smokers reported lower smoking rates while at work, although quit smoking rates and home smoking rates were similar in both hospitals. (*Am J Public Health* 1989; 79:192–195.)

Introduction

The Surgeon General's Report on involuntary smoking concludes that environmental tobacco smoke is hazardous to the health of non-smokers and that employers are obligated to protect workers from risk due to exposure.[1] In spite of the fact that many restrictive smoking policies were adopted by worksites in the 1980s,[2] most evaluations consist of casual samples of the target worksite[3] and/or retrospective reports of pre-policy behavior and attitudes.[4,5] The research reported here examined the impact over 12 months of a restrictive smoking policy adopted at one general hospital relative to a comparison hospital where no restrictive policy was implemented.

Methods

The Settings

Both institutions in this study, the Miriam Hospital (policy hospital) and the Roger Williams General Hospital (comparison hospital), are university-affiliated general medical facilities with approximately 250 beds and 1,400 employees. Before the initiation of the policy change, both hospitals had several designated no-smoking areas, and smoking was permitted in most offices and staff lounges as well as in the smoking section of the cafeteria. Patients and visitors could smoke in designated lounges on each floor and in patient rooms as long as fire safety rules were not violated.

Eight months prior to the implementation of the restrictive policy, the Employee Health Nurse brought the issue to the Safety Committee of the Miriam Hospital. During the following five months a subcommittee solicited feedback from department heads. The subcommittee eventually proposed that smoking by employees and visitors be prohibited throughout the building except for two designated sections of the hospital cafeteria and in the coffee shop. Patients would be permitted to smoke in their rooms if their roommate did not object, but smoking would no longer be permitted in patient lounges. In the interim, a new hospital president was installed who strongly supported the restrictive policy. He mandated that it be announced by July 1 and take effect on August 1, 1985.

In July a notice was included in all employees' paychecks that presented the health rationale for the policy, and stated that it would be enforced by managers and department heads in the same manner as any other issue relating to appropriate employee behavior. It also stated that free quit smoking programs would be offered on-site to employees who would like help in stopping smoking. Self-help smoking cessation programs were offered at both the policy and the comparison hospitals in the fall of 1985 as part of a larger study to be reported elsewhere.

Sampling Design

Lists were obtained of all full-time employees of the two hospitals prior to the implementation of the smoking restrictions at the policy hospital. Random cross-sectional samples were drawn without replacement from each employee list at the three assessment points: one month prior to the policy change, six months after the policy change, and 12 months after the policy change. The intention was to sample 85 employees at each hospital at each time period, under the assumption that this would yield approximately 25 smokers per sample. However, since the six-month assessment yielded only 14 smokers at the policy hospital and 18 at the comparison hospital, the sampling frame was increased to 120 employees for the 12-month assessment. The additional employees were selected from the same list of those employed at the pre-policy period. In addition to the six cross-sectional samples, a cohort of 85 individuals at each hospital was selected to be followed over the three assessment periods. There was substantial attrition from the cohort due to turnover in the labor force, vacations, and refusals. Only 71 per cent of the policy hospital cohort and 63 per cent of the comparison cohort could be re-interviewed at the final assessment period. At the policy hospital, smokers were more likely than non-smokers to refuse to be re-interviewed (seven of the 23 smokers contacted refused, versus one of the 45 non-smokers contacted; 95 per cent confidence interval for the difference between the two proportions: .06, .51). At the comparison hospital, the rate of refusal was unrelated to smoking status. Although trends for the results in the cohort were similar to those reported for the cross-sectional sample, biased attrition of the smokers in the policy hospital render their data uninterpretable. Therefore only data from the cross-sectional sample are presented in this report.

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Data Collection and Analysis

Four research assistants, trained and supervised by the first author, conducted telephone interviews with employees during working hours. Questions specific to the policy change were asked only of policy hospital employees (i.e., those measuring awareness and approval of the policy, perceptions of non-compliance and perceived effects of the policy on performance.) All other questions were asked of employees of both hospitals.

Measures taken only in the policy hospital or at both hospitals at only one time period were examined as proportions. Where appropriate, differences between groups were tested by examining differences between proportions and the 95% CI for the difference.[7] Measures taken at both hospitals at the three assessment periods were analyzed with log linear analyses (when the outcome variable was categorical) and with multiple regression analysis when the outcome was continuous (i.e., number of cigarettes smoked.) In either case, the effect of the policy change was assessed by entering the following predictors simultaneously: *Hospital* (policy hospital = 1; comparison hospital = 0); *Pre/Post-Change* (pre = −2, six months = 1, 12 months = 1), *Post/Policy-Change* (pre = 0, six months = −1, 12 months = 1); and the interactions *Hospital by Pre/Post-Change* and *Hospital by Post/Policy-Change*. The main effects estimate pre-existing differences between the hospitals as well as secular changes in the community at large. If the policy change had an impact independent of those effects, it would be detected in the interaction Hospital by Pre/Post-Change. If the effect either diminished or increased with time, it would be detected in the interaction Hospital by Post/Policy-Change.

Results

Table 1 shows the number of respondents interviewed at each hospital by time period. The refusal rates ranged from 2 to 9 per cent. Respondents were virtually identical to the employee populations from which they were drawn in sex (70 per cent female); average age (38 years); and occupational status (45 per cent professional and managerial). Characteristics of the samples from the two hospitals collapsed over the three time periods are presented in Table 2.

Almost all of the respondents at the policy hospital reported awareness of the smoking restrictions and could name the places where smoking was permitted. Over 90 per cent of the smokers and two-thirds of the non-smokers thought that the restrictive policy was "a good idea." Although half the respondents reported being aware of at least one instance of non-compliance with the policy, only 12 per cent and 16 per cent (at six and 12 months, respectively) reported more than "a few" instances. When questioned at 12 months about the effect of the policy on their performance, the majority of both smokers and non-smokers felt that it had had no effect. However, none of the smokers felt that their work had improved as compared to 21 per cent of the non-smokers who stated the improved air quality helped them to concentrate better. On the other hand, none of the non-smokers felt that their performance had deteriorated as compared to 19 per cent of the smokers.

TABLE 1—Sample Sizes by Hospital, Time Period, and Smoking Status

Hospital	Smoking Status	Time Period: Pre-Implementation	Six Months Post	Twelve Months Post
Policy	Smokers	28	14	22
	Ex and Non-smokers	54	66	88
	Total	82	80	110
Comparison	Smokers	23	18	26
	Ex and Non-smokers	60	58	78
	Total	83	76	104

TABLE 2—Characteristics of Combined Cross-Sectional Samples

Characteristics	Policy Hospital (N = 272)	Comparison Hospital (N = 263)
Average Age (years)	38	38
Sex	%	%
Female	73	67
Male	27	33
Smoking Status		
Smokers	24	26
Non-smokers	76	74
Occupational Status		
Professional Managerial	45	42
Technical	21	23
Clerical	16	17
Blue Collar Service	17	19
Education		
12 years or less	24	21
13 to 15 years	26	32
16 years or more	50	47
Marital Status		
Married	55	51
Single	32	37
Divorced, Separated or Widowed	13	12

At the 12-month assessment period, non-smokers at both hospitals were asked whether they were bothered by smoke in several areas of their workplace. Five per cent of the non-smokers at the policy hospital reported being bothered by smoke in offices compared to 25 per cent at the comparison hospital (95 per cent CI for the difference: 8,32). Results were comparable for staff lounges but the proportion of non-smokers bothered by smoke in the lavatories was similar for both hospitals (18 per cent and 16 per cent). Non-smokers were also asked to indicate what they would do if they were bothered by a coworker's tobacco smoke. Responses were coded as "assertive" if they involved saying something directly to the smoker—for example, mentioning that the smoke was bothersome. Other responses were coded as "unassertive", for example, doing nothing. Approximately half of the non-smokers at each hospital indicated that they would respond assertively.

Effect of Policy on Smoking

At each assessment period, smokers at both hospitals were asked to indicate the extent to which they agreed with the statements: "My coworkers disapprove of my smoking" and "My smoking affects the health of others." As Table 3 indicates, there was little change in response to these questions as a function of time period and little difference in response patterns in the two hospitals.

Very few smokers in the sample reported having quit during the year following the policy change. Among those individuals who reported having been smokers in August of 1985, 7 per cent at the policy hospital and 11 per cent at the comparison hospital reported quitting by one year later. Table 3 shows reports of having made a serious quit attempt in the prior six months as well as reported intentions of

TABLE 3—Smokers' Attitudes and Behaviors (with 95 per cent Confidence intervals) by Hospital and Time Period

Variables	Hospital	Time Period: Pre-Implementation	Six Months Post	Twelve Months Post
Per cent reporting coworkers disapproval of smoking	Policy	54 ± 20	46 ± 31	59 ± 23
	(n)	(28)	(13)	(22)
	Comparison	45 ± 24	41 ± 27	52 ± 22
	(n)	(20)	(17)	(25)
Per cent perceiving health risk of environmental tobacco smoke	Policy	71 ± 19	85 ± 23	59 ± 23
	(n)	(28)	(13)	(22)
	Comparison	62 ± 23	59 ± 27	76 ± 19
	(n)	(21)	(17)	(25)
Per cent considering quitting in next six months	Policy	74 ± 18	91 ± 21	27 ± 21
	(n)	(27)	(11)	(22)
	Comparison	62 ± 23	53 ± 24	65 ± 22
	(n)	(21)	(15)	(23)
Per cent who made a quit attempt in prior six months	Policy	4 ± 8	14 ± 22	0 ± 2
	(n)	(28)	(14)	(22)
	Comparison	30 ± 21	22 ± 22	19 ± 17
	(n)	(23)	(18)	(26)

quitting in the next six months. The striking difference between hospitals in the baseline level of reported quit attempts is difficult to explain. Nevertheless, according to the log linear analyses the restrictive policy had no effect of consequence on reports of quit attempts. The policy change does appear to have increased intentions to quit in the short run. This trend reversed itself by 12 months after implementation. Log linear analysis using the model described above indicates that the interaction Hospital by Post/Policy-Change was significant at the .05 level.

The effect of the policy change on smoking rate was assessed by examining the number of cigarettes smoked per day while at work, before and after work, and on non-work days with the multiple regression model described above. Age, gender, and years of education were controlled. The results of these analyses demonstrated that the number of cigarettes smoked during work decreased at the policy hospital after the restrictive policy was implemented while no such change occurred at the comparison hospital [Hospital by Pre/Post-Change: b = −1.485; standard error (b) = .664; 95 per cent CI = −2.80; −0.17]. There was no apparent effect of the policy on the number of cigarettes smoked outside of work or on non-work days. The passage of time after policy implementation had no additional effect on smoking rates. Table 4 shows the unadjusted means for the three smoking rates (age, gender, and education did not account for substantial portions of variance in smoking rates).

Records of the self-help smoking program offered at both hospitals soon after the baseline assessment indicate that the policy did not lead to an increase in the enrollment of smokers. Thirteen per cent of smokers at the policy and 14 per cent at the comparison hospital enrolled in the program.

Discussion

Other studies have also reported high rates of approval for restrictive policies, even among smokers.[3,6] In only one published report did a majority of the smokers disapprove of smoking restrictions.[5] In that case the policy created a total ban on smoking and the implementation strategy apparently left many of the smokers confused about their role in the

TABLE 4—Daily Smoking Rate (with 95 per cent confidence intervals) by Hospital and Time Period

Variables	Hospital	Time Period: Pre-Implementation	Six Months Post	Twelve Months Post
Mean cigarettes smoked per day at work	Policy	8.4 ± 2.5	3.5 ± 1.5	4.5 ± 1.4
	(n)	(28)	(16)	(23)
	Comparison	7.6 ± 2.6	7.4 ± 3.1	6.9 ± 2.0
	(n)	(21)	(21)	(29)
Mean cigarettes smoked per day at home	Policy	12.8 ± 2.5	9.4 ± 3.9	10.6 ± 3.1
	(n)	(28)	(16)	(23)
	Comparison	13.3 ± 4.1	7.7 ± 3.4	9.2 ± 2.6
	(n)	(21)	(21)	(29)
Mean cigarettes smoked on non-work days	Policy	19.3 ± 2.9	13.2 ± 4.6	16.3 ± 3.6
	(n)	(28)	(16)	(23)
	Comparison	19.3 ± 5.8	12.5 ± 4.8	13.9 ± 4.1
	(n)	(21)	(21)	(29)

*Analyses of mean cigarettes smoked include scores of zero for 3 individuals at the policy hospital and 6 at the comparison hospital who reported having quit smoking since the pre-policy assessment. Unadjusted means are presented.

decision-making. Employers' concern about severe dissension between smokers and non-smokers has been identified as a major barrier to the implementation of restrictive smoking policies at the worksite. The results of the present study can be added to the mounting evidence that smoking can be restricted at the workplace without undue conflict or dissatisfaction.

The results suggest that restrictive policies do lead to a reduction in self-reported smoking behavior sustained for up to 12 months after implementation. Furthermore, smokers who reduced their smoking at work did not seem to compensate for their reductions by increasing their smoking outside of work as theories of nicotine dependence might predict.[8,9] The fact that different individuals were surveyed at each assessment requires that some caution be used in interpreting the results. However, analysis of the effects of shifts in the demographic composition of the samples (i.e., shifts in age, job status, and education) indicate that patterns of smoking and quitting cannot be attributed to these shifts. Although the present study, as well as previous work, lacks biochemical validation of reports of smoking behavior, the validity of smokers' reports is bolstered by non-smokers report of less annoyance from second-hand smoke.

The present results do not support the expectation that restrictive policies will increase the rate of smoking cessation either by making non-smokers less tolerant of smokers, or by directly affecting the motivation and/or quitting efforts of smokers. Our ability to detect effects on smoking behavior is limited by the small number of smokers in our sample, so this conclusion must be seen as tentative. However, the fact that enrollment in a free worksite-based self-help program was similar in the two hospitals can be seen as an additional indication that smokers' motivation to quit was not influenced by the policy. Furthermore, only one published study reports high rates of quitting subsequent to a restrictive policy,[3] and that finding is limited by the non-representativeness of the sample and the lack of a control group. The fact that intentions to quit increased shortly after the policy was implemented, and then decreased to below baseline levels may indicate that smokers expect that restrictive policies will require them to quit. When they learn that they can adapt, by going to the remaining smoking areas and by reducing the number of cigarettes smoked at work, the motivation to quit may be undermined. If so, the most opportune time for a quit smoking campaign may be during the early phases of a policy change.

It has been expected that restrictive policies will mobilize non-smoker pressure on smokers, yet 50 per cent of the non-smokers report being reluctant to ask a co-worker to refrain from smoking in their presence. Furthermore, to the extent that a policy is successful in reducing non-smokers' exposure to smoke, they may find even less need to put pressure on smokers and this potential source of motivation to quit may decrease.

Research to date suggests that no single intervention in isolation is sufficiently powerful to treat a severe addiction like smoking. Dramatic reductions in smoking prevalence may require multifaceted interventions including environmental restrictions, incentives to quit and treatment programs targeted to specific subgroups.[10] Nevertheless, reductions in rate of smoking over 12 months is a promising result of restrictive policies, and worksites still have potential for contributing to the goal of a smoke-free society.

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Nominations Invited for Spiegelman Award in Health Statistics

Nominations are invited for the 1989 Spiegelman Gold Medal Award. This annual award is given to a candidate born in 1950 or later who has made substantial contribution to the practice or theory of health statistics, broadly defined. Nomination forms may be obtained from Ms. Kinaya Sokoya, American Public Health Association, 1015 Fifteenth Street, NW, Washington, DC 20005, and must be returned to her by March 31, 1989.

Public Health Briefs

Cigarette Smoking among San Francisco Hispanics: The Role of Acculturation and Gender

Gerardo Marin, PhD, Eliseo J. Perez-Stable, MD, and Barbara Vanoss Marin, PhD

Abstract: We conducted a phone survey of 1,669 San Francisco Hispanics ages 15 to 64 years. The age-adjusted overall smoking prevalence was 25.4 per cent (95% CI = 23.3, 27.5) with more men (32.4 per cent) smoking than women (16.8 per cent). Age-adjusted smoking rates were higher among the less acculturated males (37.5 vs 26.7 per cent) and among the more acculturated females (22.6 vs 13.6 per cent). The more acculturated, however, smoked a greater number of cigarettes per day independent of gender. Community-based smoking cessation interventions, adapted to local conditions, may have a greater potential for success among Hispanics. (*Am J Public Health* 1989; 79:196–199.)

Introduction

Most studies on cigarette smoking among Hispanics indicate that although the overall prevalence is lower than the rates for non-Hispanics Whites, Hispanic men smoke more than Hispanic women.[1–7] The level of acculturation to the mainstream US culture may play a significant role in influencing cigarette smoking behavior. However, the relationship between acculturation and smoking prevalence remains unclear, in part due to the fact that instruments used to measure acculturation have usually been plagued by serious psychometric shortcomings.[8] The southwest portion of the Hispanic Health and Nutrition Examination Survey (HHANES)* showed that 28 per cent of women scoring in the upper tertile of an acculturation scale reported current smoking, compared to 19 per cent of those scoring in the lower tertile of acculturation. Among the men sampled in the HHANES, the level of acculturation was not consistently related to cigarette smoking. In contrast, a recent study among a representative sample of older Mexican Americans and of their relatives residing in San Antonio, Texas showed that acculturation was not related to prevalence of smoking.[3]

To evaluate the role of acculturation in cigarette smoking behavior, we conducted a phone survey of a representative sample of Hispanics in San Francisco.

Methods

Respondents were sampled using Waksberg's[9] method for random digit dialing from the 27 census tracts in San Francisco with 10 per cent or more Hispanic residents. These tracts represented 66.6 per cent (N = 55,541) of all Hispanics (N = 83,373) living in the city of San Francisco according to the 1980 census.[10] Participants were given a choice of answering the interview in either English or Spanish; 67 per cent preferred to be interviewed in Spanish. The refusal rate (proportion of eligible respondents contacted who declined to be interviewed) was only 4.7 per cent. Hispanic ethnicity was defined by self-identification.

The interview schedule asked about current and past smoking behavior, awareness of smoking cessation programs, standard sociodemographic items, and a five-item acculturation scale developed by Marin and colleagues which has shown excellent reliability and validity.[8]

Data were analyzed using standard techniques. Age-adjusted rates were calculated by direct standardization method[11] using the 1980 Hispanic population in San Francisco as reference.

Results

A total of 1,669 Hispanics, between 15 and 64 years of age, were interviewed by telephone between October 1986 and March 1987. The respondents' had a mean age of 35.5 years (SD = 12.06), an average of 11.6 years of education and had lived in the United States a mean of 14.2 years; 69.5 per cent were born in Latin America. Interviewed Hispanics were of Mexican (36.4 per cent), Central American (46.9 per cent), South American (6.1 per cent), or Caribbean (10.6 per cent) background.

A total of 735 (44 per cent) respondents (95% confidence intervals = 41.6, 46.4) reported smoking at least 100 cigarettes in their lives. The overall age-adjusted prevalence of current smoking (Table 1) was 25.4 per cent (95% CI = 23.3, 27.5) but only 16.8 per cent (95% CI = 15.6, 18.0) of the women compared to 32.4 per cent (95% CI = 29.4, 35.4) of the men were currently smoking. Less acculturated men had a higher age-adjusted smoking rate than the more acculturated (37.5 vs 26.7 per cent), while the reverse relation held with women (13.6 vs 22.6 per cent).

The more highly acculturated Hispanic smokers reported smoking an average of 13.3 cigarettes per day compared to 8.6 cigarettes per day among the less acculturated and waiting an average of over 90 minutes less than those with

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*Haynes SG, Cohen B, Harvey C, McMillen M: Cigarette smoking patterns among Mexican-Americans: HHANES, Southwest United States, 1983-1984. Paper presented at the Annual Convention of the American Public Health Association, Washington, DC, November 1985.

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TABLE 1—Smoking Behavior among San Francisco Hispanics, Ages 15 to 64 Years

	Total N	Current Smokers* N	Current Smokers* %	Former Smokers N	Former Smokers %	Never Smokers N	Never Smokers %
Total	1669	406	25.4	329	19.7	934	56
Men	731	245	32.4	192	26.3	294	40.2
Low Acculturation	435	167	37.5	106	24.4	162	37.2
High Acculturation	295	78	26.7	86	29.2	131	44.4
Women	938	161	16.8	137	14.6	640	68.2
Low Acculturation	623	88	13.6	78	12.5	457	73.4
High Acculturation	313	72	22.6	59	18.8	182	58.1

Total per cents may not add up to 100 due to rounding.
*Rates for current smoking are age-adjusted using the 1980 Hispanic population in San Francisco as reference.

lower levels of acculturation before lighting their first cigarette of the day (Table 2).

The less acculturated respondents tended to report quitting plans to a greater extent than the more acculturated (Table 3). For example, the mean number of quit attempts reported for the 12 months prior to the interview was 2.0 for the less acculturated compared to 1.4 for the more acculturated. Less acculturated male smokers also reported 22 more days without smoking in the previous year, but among women there was virtually no difference. The highly acculturated felt more certain than the less acculturated about how to obtain information dealing with smoking cessation and about where to join a cessation group. However, all Hispanic smokers interviewed reported having heard information on the health dangers of smoking in a similar proportion.

Discussion

The results of this survey of San Francisco Hispanics are consistent with other published reports that have shown marked differences in smoking prevalence due to the gender and acculturation level of the respondents.[1-7] Our data also indicate a lower smoking prevalence than found in regional surveys[3,6,*] completed several years ago. This may actually represent a trend toward decreasing smoking prevalence

TABLE 2—Smoking Behavior of Current Smokers by Gender and Acculturation

	Men (N = 245)		Women (N = 161)		Total (N = 406)	
	Acculturation		Acculturation		Acculturation	
(N)	Low (167)	High (78)	Low (88)	High (72)	Low (255)	High (150)
Mean number of cigarettes smoked/day	8.6	15.6	6.8	10.8	8.0	13.3
Mean age started smoking	16.7	16.6	21.1	17.2	18.2	16.9
% Inhaling deeply into chest	48	68	30	54	42	61
% burning all cigarette	38	35	15	35	31	35
Average wait (min) to first cigarette upon waking	272	178	328	219	291	198

Note: All percentages rounded to nearest whole number.

TABLE 3—Quitting Behavior and Awareness of Cessation Services among Current Smokers by Gender and Acculturation

	Men (N = 245)		Women (N = 161)		Total (N = 406)	
	Acculturation		Acculturation		Acculturation	
(N)	Low (167)	High (78)	Low (88)	High (72)	Low (255)	High (150)
Plan to change smoking patterns next year	71%	62%	85%	54%	75%	58%
Mean number of quit attempts last year	2.0	1.4	2.1	1.4	2.0	1.4
Mean number of days without smoking last year	62.9	39.1	46.8	50.1	57.5	44.0
% Knows of printed cessation information	26	57	39	65	31	61
% Knows of cessation group	14	49	15	46	14	48
% Heard/Read information on health hazards	81	73	80	70	81	71

Note: Percentages rounded to nearest whole number.

among Hispanics since our rates (32.4 per cent for men and 16.8 per cent for women) are similar to those reported for Hispanics in the 1985 National Health Interview Survey[7] (31 per cent and 21 per cent, respectively).

However, several limitations of this study must be considered. First, although the majority of San Francisco Hispanics live in the area included in the survey, the sample may not be representative due to the 66.6 per cent coverage. Second, deception in self-reporting cigarette smoking behavior has been reported among New Mexico Hispanics,[12] and thus the low smoking prevalence found here may be an underestimate. Third, Hispanics in San Francisco live in a strong anti-smoking environment and thus may be influenced to an extent that limits generalizability to Hispanics in other areas. Finally, the acculturation scale applied is a unidimensional measure of a complex phenomenon that is otherwise difficult and tedious to evaluate.

Despite these limitations, the observed differences in prevalence and smoking behavior by the respondents' level of acculturation are consistent with the concept that as individuals adapt to a new culture their attitudes and behaviors also change.[13] These data indicate that the influence of acculturation on smoking prevalence and behaviors varies by gender. Our observation that increased acculturation among men results in decreased smoking, and that the opposite effect is seen among women, may reflect current social norms for smoking in the United States. In addition, the number of cigarettes smoked per day and other smoking behaviors of the more acculturated Hispanics resemble those of non-Hispanic White smokers.

Future prevention and cessation programs directed at Hispanics will need to consider the influence of acculturation on the prevalence and smoking behaviors of prospective clients. In addition, the less acculturated, primarily Spanish-speaking Hispanic seems to be in greater need of information and services. A majority of Hispanics report being light smokers[1,3,*] implying that less intense cessation methods may be expected to have a greater impact among them. Thus large-scale community programs utilizing Spanish-language mass media with information on the dangers of smoking and including simple

cessation techniques that smokers with low levels of addiction can utilize to quit smoking need to be a priority. The need to target the less acculturated Hispanic men, as the subgroup with the highest prevalence, is a challenge to program planners who will need to deal with stereotyped beliefs regarding smoking among the members of this group.[14]

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Evidence for Limits on the Acceptability of Lowest-Tar Cigarettes

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Abstract: The sales of the lowest yield cigarettes (1-3 mg tar) seem to have been particularly resistant to the effects of promotion and advertising, while the sales of other low-yield cigarettes (4-9 mg tar) seem to have been increased by promotional efforts. This finding is consistent with the existence of a boundary of tar and nicotine acceptability below which consumers in general are not prepared to go. Use of lower tar cigarettes may be helpful for those who cannot stop smoking, but, since 1979, the percentage of cigarettes under 16 mg tar has changed little. (*Am J Public Health* 1989; 79:198–199.)

Introduction

There may be limits to consumer acceptability of the lowest-yield cigarettes.[1,2] Although these cigarettes have been promoted (sometimes as less hazardous), they have attracted few customers.[2] From information provided by the cigarette industry, the United States Federal Trade Commission (FTC) has reported the promotional costs and sales of cigarettes as a function of tar yield categories for the years 1979 to 1986[3,4] (see Table 1). The relation between promotional costs and sales is least ambiguous when promotional costs are increasing: Other things being equal, greater promotion should be reflected in greater sales. When promotional costs drop, brand loyalty and other habitual pressures can account for much of the repeat purchasing. I therefore assessed the relation between promotion and sales for the three consecutive years in which promotional costs were increasing (1979 through 1981) for each of the three lowest-yield categories of low-tar cigarettes.

Method

Advertising and promotional expenditures include advertising in newspapers, magazines, paperback books, outdoor billboards, transit posters, and at point of sale, as well as promotional allowances, sampling and other distribution, sponsorship of special events, and costs relating to discount coupons.[3,4] A linear regression was calculated between the percentage of the advertising and promotional budget and the percentage of sales (market share) for each of three tar yield categories (3 mg tar or less, 4–6 mg tar, and 7–9 mg tar). The

TABLE 1—Percentage of Promotional Advertising (Adv) Costs and Sales for Cigarettes in Three Categories of Tar Yield and Total Adv Costs and Sales from 1979 to 1986

	1–3 mg Tar		4–6 mg Tar		7–9 mg Tar		Totals	
Year	Adv	Sales	Adv	Sales	Adv	Sales	Adv (in millions)	Sales (in billions)
	%	%	%	%	%	%	$	$
1979	9.6	2.7	9.3	3.1	12.0	4.8	1,083	621.8
1980	15.1	3.3	9.9	4.0	14.7	9.5	1,242	628.2
1981	15.7	3.7	11.5	5.9	19.5	15.0	1,548	636.5
1982	5.7	2.9	9.9	6.0	25.8	18.9	1,794	632.5
1983	4.2	3.1	11.5	6.3	19.4	18.5	1,901	603.6
1984	4.3	2.9	8.0	6.5	21.1	16.9	2,095	608.4
1985	3.1	2.3	6.4	6.1	20.6	16.9	2,476	599.3
1986	3.8	2.6	7.7	7.3	14.6	12.4	2,382	586.4

SOURCE: US Federal Trade Commission Reports to Congress[3,4]

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regression analyses were limited to the years 1979 through 1981 because these were years in which promotional budgets were increasing each year for the three yield categories. Detailed data were not available from earlier years. Other patterns in the data in Table 1 are discussed.

A 0.4 to 0.5 per cent market share of this huge market is generally considered successful.[5,6] In 1979, a 1.0 per cent market share could supply just over 850,000 one pack-a-day smokers with cigarettes for one year; and a 1.0 per cent share of the promotional budget was worth $10.8 million. A difference of 1 percentage point in market share or promotional budget was considered substantial and worth discussing.

Results and Discussion

The regression results for 4–6 and 7–9 mg tar were so similar that a single regression is reported for the 4–9 category. Sales of 4–6 and 7–9 mg tar cigarettes were strongly related to the promotional expenditures (slope = 1.17, SE = .07), while the sales of 1–3 mg tar cigarettes were only weakly related to promotional costs (slope = 0.14, SE = .047) (see Figure 1). This supports the boundary model of smoking which argues that smokers may find a wide range of nicotine doses acceptable, provided that a minimum, "lower boundary" level of nicotine intake can be maintained.[1,7]

After the increase in promotion from 1979 to 1981, there was a leveling off for the 7–9 mg tar cigarettes, and a substantial drop in promotion of both the 4–6 and the 1–3 mg tar cigarettes (14 per cent and 64 per cent decrease); sales remained steady for the 4–6 mg tar cigarettes, but dropped for the 1–3 mg tar cigarettes (2 per cent increase vs 22 per cent decrease) (see Table 1). (The 1982 values for sales and advertising for 7–9 mg tar cigarettes do fall on the same regression line as for the earlier years.) Most of the sales of the 1–3 mg tar cigarettes may have been due to smokers temporarily trying these cigarettes rather than adopting them as their brand; hence when the promotion declined, the trial purchases declined. The 4–6 mg tar cigarettes may have been more likely to find consumers who adopted them as their brand, hence the steady sales when the promotion declined. There is also evidence for a sensible "tuning" of promotional costs and sales over the years. Those interested in making a profit cannot be expected to continue promoting products that are not selling. The historical pattern suggests that with promotion consumers can be drawn to brands in the 7–9 and even the 4–6 mg tar ranges, but this may be increasingly difficult for the kind of individuals (i.e., lower socioeconomic status, possibly more addicted) who remain as smokers in spite of years of pressure against smoking.[8]

In 1977, some health authorities thought that anti-cigarette legislation would be forthcoming to limit tar yields to a maximum of 15 mg (the FTC designation for "low-tar")[9]; and the president of one cigarette company suggested that market forces rather than legislation would "soon" bring about a 15 mg tar limit.[9] Currently, a 15 mg tar limit is far from a reality. The sales-weighted trend toward smoking lower and lower yield cigarettes has been flat since 1982.[10] In 1986, 47.4 per cent of the cigarettes sold in the United States delivered at least 16 mg tar.[4] Although there is dispute about the extent to which lower yield cigarettes are less hazardous than higher yield cigarettes, there are few who would argue that higher yield cigarettes are less hazardous.[11] Although the health benefits of such a market change are likely to be very small to individual smokers, they may yield worthwhile public health benefits.[11,12]

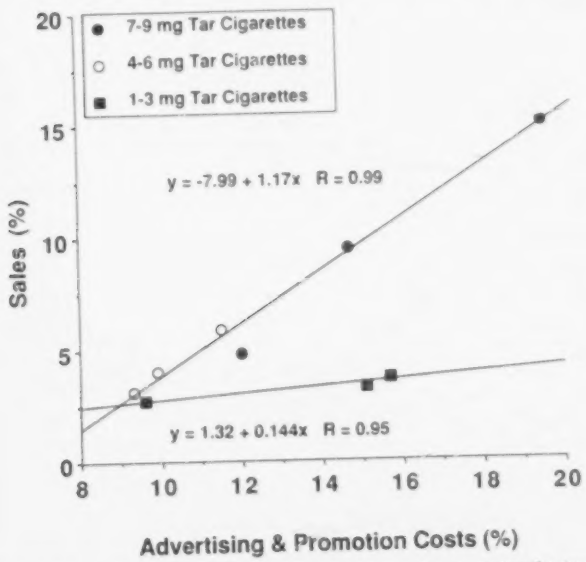

FIGURE 1—Percentage of sales of cigarettes in three low-tar categories (see insert) as a function of percentage of promotional/advertising budget for the years 1979, 1980, and 1981. For each tar category, the left-most point is 1979 and the right-most point is 1981. (Note that the abcissa is plotted from "8" and not zero.)

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Accuracy of Recalled Smoking Data

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Abstract: Validity of recalled cigarette data was assessed among 87 middle-aged adults by comparing recall to longitudinal records. Agreement on smoking status and amount smoked 20 years ago occurred for 87 per cent and 71 per cent of subjects, respectively. Corresponding proportions for 32-year recall were 84 per cent and 55 per cent. Gender did not influence recall accuracy. Ex-smokers tended to make the most errors. Accuracy of recalled smoking information 20 years ago was comparable to that of alcohol status and consumption frequency category. (*Am J Public Health* 1989; 79:200–206.)

Introduction

Information obtained by recall is vulnerable to numerous inaccuracies and biases; nevertheless, retrospective methods are commonly used to determine past exposure to such substances as tobacco and alcohol. Studies evaluating the accuracy of long-term recall (10 years or more) of smoking data are few.[1–3] Although faulty recall of prior smoking status was rare, more than half the respondents in one study incorrectly reported amounts smoked 15 years previously.[1] Exact agreement between recall of alcohol consumption frequency and that recorded 15 years before occurred for approximately half the respondents in a community study.[4] Overestimation of alcohol consumption has been noted.[5]

This paper examines the accuracy of retrospective smoking reports at two time points (20 and 32 years earlier) among a group of middle-aged men and women. Recall of smoking 20 years ago is compared to that of alcohol consumption at the same time point.

Methods

The participants were enrolled prenatally between the years 1930 and 1938 in a longitudinal study of health and development.[6] The original cohort (n=134) was drawn from middle-class, Caucasian families living in Boston. During childhood (until age 18), subjects were examined and interviewed annually. Further contact occurred in 1965 and 1984–85. At age 18 and 1965 interviews, smoking and dietary histories were obtained by use of standardized questionnaires.

Forty-three men and 52 women, the majority of whom resided in New England, participated in the 1984–85 follow-up. The mean educational level was 14 years, and ranged from 10th grade to professional degrees. Age ranged from 47 to 54 years.

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Subjects completed self-administered questionnaires on which they recalled smoking status, number of cigarettes or packs per day, and consumption frequencies of beer, wine, and liquor approximately 32 years before (at age 18 follow-up) and 20 years before (in 1965), in that order. Recall calendars helped them focus on the appropriate time points. Current smoking status and amount smoked were also obtained. Two subjects who failed to finish the 1984–85 interview and six men who smoked only pipes or cigars were subsequently excluded from analyses.

Validity of retrospective smoking and alcohol reports was assessed by comparing them to data recorded at previous interviews. Actual and recalled cigarette doses were coded into four categories: none; <1 pack/day; 1–2 packs/day; or ≥2 packs/day. Conversion to categorical data was done because fractions of packs rather than number of cigarettes were most often reported. Actual and recalled weekly consumption frequencies of beer, wine, and liquor were summed and the totals recoded into four categories: none, <1 drink/day; 1–3 drinks/day; or ≥4 drinks/day. Proportions of subjects recalling an amount in a category consistent with recorded data were evaluated by kappa statistics[7] and differences between proportions by chi-square statistics.

Current smoking status definitions were based on 1984–85 reports. Current nonsmokers were subclassified into those who were nonsmokers both 20 and 32 years ago, and those who smoked at either earlier time point.

Results

The proportion of smokers in this group decreased from a maximum of 59 per cent 20 years earlier to 33 per cent at present; half the current smokers had reduced the number of cigarettes per day since they began smoking. In contrast, the proportion of alcohol users increased by 14 per cent during the same interval and mean consumption rose by three drinks per week.

Table 1 presents the agreement between recalled and historically recorded smoking status 20 and 32 years ago. The kappa statistics indicate better than chance agreement. There were no differences in accuracy by gender, current smoking status, or length of recall interval.

Amount smoked was recalled less accurately than smoking status (Table 2). Current smoking status, but not gender, influenced recall accuracy. A significantly smaller proportion of ex-smokers than both never-smokers or current smokers correctly recalled cigarette smoking level 20 years ago and 32 years ago. Even though recall was better than chance, a substantial proportion were misclassified by their retrospective reports. Further examination of these reports revealed that underestimation of amount (17 per cent for the 20-year and 37 per cent for the 32-year recall) was more common than overestimation (9 per cent and 6 per cent).

The proportions of respondents correctly reporting past smoking status (90 per cent) and cigarette dose (74 per cent) 20 years earlier were similar to those correctly reporting alcohol use status (86 per cent) and consumption frequency (61 per cent). Men and women reported equally well. Unlike smoking, however, underreports of alcohol consumption

TABLE 1—Recall of Cigarette Smoking Status 20 and 32 Years Previously by Sex and Current Smoking Status

Respondents	n	Number and % of Individuals Who Correctly Recalled Smoking Status: 20 years ago, n (%) correct	20 years ago, kappa[a]	32 years ago, n (%) correct	32 years ago, kappa[a]
Total	87	78 (90)	.80	75 (86)	.72
Men	36	30 (83)	.65	32 (89)	.74
Women	51	48 (94)	.88	43 (84)	.66
Current Smoker[b]	29	28 (97)	.81	27 (93)	.85
Men	10	9 (90)		9 (90)	
Women	19	19 (100)		18 (95)	
Current Nonsmoker	58	50 (86)	.71	48 (83)	.66
Men	26	21 (81)		23 (88)	
Women	32	29 (86)		25 (78)	
Current Nonsmoker Past Nonsmoker[c]	29	27 (93)		24 (83)	
Men	8	7 (88)		7 (88)	
Women	21	20 (95)		17 (81)	
Current Nonsmoker Past Smoker	29	23 (79)		24 (83)	
Men	18	14 (78)		16 (89)	
Women	11	9 (82)		8 (73)	

a) All kappas significantly greater than 0, $p < 0.01$.
b) All but one individual were also past cigarette smokers.
c) Reported non-smoking status for both past time periods.

TABLE 2—Recall of Cigarette Consumption Level 20 and 32 Years Previously by Sex and Current Smoking Status

Respondents	n	Number and % of Individuals Who Correctly Recalled Smoking Dose[a]: 20 years ago, n (%) correct	20 years ago, kappa[b]	32 years ago, n (%) correct	32 years ago, kappa[b]
Total	87	64 (74)	.63	50 (57)	.36
Men	36	23 (64)	.50	17 (47)	.31
Women	51	41 (80)	.71	33 (65)	.39
Current Smoker	29	21 (72)	.61	15 (52)	.36
Men	10	5 (50)		2 (20)	
Women	19	16 (84)		13 (68)	
Current Nonsmoker	58	43 (74)	.56	35 (60)	.33
Men	26	18 (69)		15 (58)	
Women	32	25 (78)		20 (63)	
Current Nonsmoker Past Nonsmoker	29	27 (93)		24 (83)	
Men	8	7 (88)		7 (88)	
Women	21	20 (95)		17 (81)	
Current Nonsmoker Past Smoker	29	16 (55)		11 (38)	
Men	18	11 (61)		8 (44)	
Women	11	5 (45)		3 (27)	

a) Categories: 1) nonsmoker
2) <1 pack (1–19 cigarettes) day
3) 1–2 packs (20–39 cigarettes) day
4) ≥2 packs (≥40 cigarettes) day
b) All kappas significantly greater than 0, $p < 0.01$.

level (1½ per cent) were less common than overreports (28 per cent).

Discussion

In the present investigation, we extended the length of recall of cigarette use beyond the 10–18 years reported by others.[1–3] Our findings are consistent with those previously cited. Recall of past smoking status was valid, but poor memory of quantitative detail such as cigarette dose, which is often of greater interest, led to frequent misclassification of prior consumption level, particularly by ex-smokers. The somewhat better recall of cigarette dose by these individuals compared to a previous study[1] may have been due to a higher proportion of never-smokers, whose recall was most accurate, or an effect of the life-long attention to and reporting of health-related information required of this cohort. Smoking and alcohol habits were remembered equally well.

Our data are also consistent with observations that current smoking or alcohol levels bias estimates of past use.[1,4,5] Present smoking levels were lower than those 20 years ago and retrospective estimates understated true past cigarette use. Overestimation of alcohol consumption by the recall method was observed in conjunction with an increased

prevalence of this practice. Alternatively, differences in social acceptability of the two practices may have influenced recall.

A limitation of this study and others with similar design[1–3] that makes it difficult to generalize to any smoking recall task is that subjects were asked to remember smoking habits at brief and specific time points. Continuous practices may be easier to remember, possibly making comprehensive smoking histories more accurate. This issue could be better resolved by administering retrospective smoking questionnaires to individuals whose smoking habits have been recorded prospectively at frequent intervals.

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Validity of Telephone Surveys in Assessing Cigarette Smoking in Young Adults

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Abstract: Self-reports of cigarette smoking in a sample of young adults, ages 17 to 21 (n = 359) obtained by telephone were validated by home interview and saliva cotinine measurement. While most subjects were in agreement for their reported behaviors, the telephone methodology underestimated cigarette smoking rates by 3 to 4 per cent. Reported quitting by telephone was an unstable category with 35 per cent of quitters reporting smoking on home interview. Reliance on telephone methods alone may underestimate smoking behavior. Validation of surveys, preferably by biochemical methods, is recommended. (*Am J Public Health* 1989; 79:202–204.)

Introduction

Telephone interviews have certain advantages relative to cost, convenience, and participation compared to face-to-face interviews. However, it is legitimate to question telephone methods for determining behaviors such as cigarette smoking, particularly in populations known to be likely to underreport—e.g., adolescents,[1] participants in cessation programs.[2] Information comparing telephone to face-to-face interview methods is limited but the results have not been reassuring. One repeatability study reported poor comparability in the number of cigarettes smoked per day, the age of quitting, and the age of initiating smoking among young Australian veterans (ages 29–39).[3] Massey, *et al*,* compared telephone and home interview methods finding 17–19 year-old males had a tendency in telephone interviews to underreport their current smoking and overreport quitting. Neither of these studies used biochemical measures to validate the smoking self-reports.

To better assess the reliability and validity of the telephone interview for cigarette smoking behavior among young adults, comparison was made to subsequent face-to-face interviews validated by saliva cotinine.[4,5] The results and the implications for survey data gathered by telephone methods are presented.

Methods

Subjects were young adults (ages 17 to 21) living in the Minneapolis-St. Paul metropolitan area and representing the classes of 1985 and 1986 from seven area high schools. They were interviewed by telephone between July 1, 1986 and June 30, 1987 to collect information on smoking attitudes and behavior. Participants were representative of a cohort of 7,124 adolescents followed annually to assess changes in smoking habits during young adulthood.[6]

Four groups (described below) were defined: non-smokers, smokers, and short- or long-term quitters. The sampling fraction for each group was based on estimates of their prevalence in the whole population to provide adequate sample size to detect meaningful differences between telephone and home interview responses: 14 per cent non-smokers, 20 per cent smokers, and 81 per cent quitters who resided in the metropolitan area. Sampling was random within each group. Subjects were invited for a home interview 11 to 31 days after telephone interview. Participants were paid $20 to answer an interviewer-administered questionnaire including the tobacco use questions asked by

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telephone and to provide a saliva sample for cotinine analysis by gas chromatography.[7] Methods of saliva sample collection are described elsewhere.[8]

The four groups according to their cigarette smoking status at the phone interview were: *Non-smokers* were those who did not smoke at the time of the interview, had smoked less than 100 cigarettes during their lifetime, or had quit smoking more than one year previously. *Smokers* were those currently smoking who had smoked more than 100 cigarettes in their lifetime. *Short-term quitters* had stopped smoking during the 60 days before the telephone interview. *Long-term quitters* had quit smoking more than 60 days but less than 365 days.

Results

Of 6,712 subjects surveyed by phone (participation = 94.2 per cent; 6,712/7,124), 729 were asked to participate in the home interview, and 587 complied (80 per cent). As shown in Table 1, non-smokers not in school and male quitters were somewhat less likely to participate in the home interview. This may be due partly to the requirement of still living in the metropolitan area; males and university students are more likely to out-migrate. The data presented here include only 359 of the 587; the difference is due to inaccurate cotinine measurements caused by laboratory equipment failure during the first part of the study. Subjects with these samples did not differ from those in the analyses in reported smoking habits.

Comparison of the telephone interview with the home interview (Table 2) shows that agreement was high among non-smokers (95 per cent), and smokers (97 per cent). The lowest agreement was found among those calling themselves short-term quitters during telephone interview. On the home interview, many subsequently reported smoking (59 per cent). A smaller but substantial proportion of those reporting by telephone to be long-term quitters also subsequently reported smoking on the home interview (28 per cent).

Table 3 compares self-reported smoking by telephone and home interview with saliva cotinine. Previous studies in this laboratory with known smokers and non-smokers suggest that 20 ng/ml provides good sensitivity and specificity for detecting regular cigarette use in this age group (data available from first author). Individuals who used smokeless tobacco one week prior to the home interview were excluded. Compared to the telephone interview, most reported non-smokers (96 per cent) had cotinine levels less than 20 ng/ml while most smokers (93 per cent) had cotinines above 20 ng/ml. Short-term quitters, by telephone interview, were likely to have elevated cotinine levels (54 per cent), as were reported long-term quitters (27 per cent).

TABLE 1—Demographic Characteristics of the Home Visit Participants and the Total Sample

Non-smokers		Smokers		Quitters	
Home	All	Home	All	Home	All
Average Age (years)					
19.00	18.94	19.01	18.93	18.96	18.98
Gender (% Male)					
50.56	52.68	44.10	47.50	45.71	51.05
% Graduated from High School					
97.50	96.44	80.51	75.15	88.00	85.25
% Current Student					
45.83	55.45	30.77	28.05	38.86	41.22

TABLE 2—Smoking History: Telephone vs Home Interview

	Home interview				
Telephone Interview	Non-smoker	Smoker	Short-term Quitter	Long-term Quitter	Total
Non-smoker	175	5	1	4	185
Smoker	1	87	2	0	90
Short-term Quitter	1	10	4	2	17
Long-term Quitter	9	19	2	37	67
Total	186	121	9	43	359

The home interview also showed high concordance between cotinine and smoking status in smokers and non-smokers; however, the home interview also showed higher concordance with cotinine for short-term quitters (100 per cent) and long-term quitters (93 per cent).

Table 4 extrapolates the marginal prevalence rates presented in Table 2 to the entire cohort of 6,712 subjects based on the sampling fractions used in this study. As is apparent, prevalence based on the home interview suggests a higher rate of smoking and a lower rate of non-smoking than do the rates based on the telephone interview; the prevalence of quitting is relatively unaffected by the source of the data. McNemar's test for symmetry confirmed that these differences were greater than would be expected by chance.

Discussion

These data suggest that telephone survey methods underestimate smoking rates and overestimate non-smoking. Both long- and short-term quitting by telephone interview, while constant in size, are unstable categories with subjects leaving and entering in the short interval between surveys. Extrapolating the home survey data to the entire population as seen in Table 4 increases smoking rates by 3.1 per cent (24.2 to 27.3 per cent). A portion of this derives from a few individuals who described themselves as non-smokers on the telephone but reported smoking at home interview. These are likely deceivers. The remainder of the smoking increase results from short- and long-term quitters on the telephone who reported regular smoking at home interview. This may be due to real relapse but may also be underreporting on the telephone.

This study observed a constant proportion of short- and long-term quitters. However, this category was highly unstable with a substantial portion of long- and short-term quitters on the telephone reporting smoking on home inter-

TABLE 3—Distribution of Self-reported Smoking at the Telephone and Home Interviews as a Function of Cotinine Level

	Smoking History: Telephone vs Saliva Cotinine							
	Non-smoker		Smoker		Short-term Quitter		Long-term Quitter	
Cotinine Cutpoint (mg/ml)	N	(%)	N	(%)	N	(%)	N	(%)
Telephone Interview								
<20	169	(96)	6	(7)	6	(46)	45	(73)
≥20	7	(4)	81	(93)	7	(54)	17	(27)
Home Interview								
<20	170	(97)	12	(10)	6	(100)	38	(93)
≥20	5	(3)	103	(90)	0	(0)	3	(7)

*Smokeless tobacco users in past week excluded.

TABLE 4—Population Prevalence Rates (per hundred subjects) Based on Extrapolation of the Marginal Rates in Table 2

	Non-smoker	Smoker	Quitter
Telephone interview	70.1	24.2	5.7
Home interview	67.3	27.3	5.5

view. This underreporting of smoking was balanced by individuals who reported non-smoking on the telephone but describe themselves as quitters at home.

The use of cotinine as a validating tool has two effects. First, chemical measurement tends to increase accurate reporting of smoking on interview.[1] In addition, it can appropriately classify individuals who continue to underreport even on direct interview. The use of cotinine measurement adds further to the smoker category. Smoking prevalence, as defined by cotinine, would be 28.6 per cent, or a 1.3 per cent increase over the home interview level.

In summary, telephone interviews for smoking behavior may lead to 3–4 per cent underreporting of smoking behavior. Quitting is an unstable category because of relapse and misreporting. These factors need to be considered when performing telephone interviews for smoking or other proscribed behaviors. Wherever possible, validation, preferably by biochemical measures, should be included to accurately estimate prevalence.

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Patient Literacy and the Readability of Smoking Education Literature

CATHY D. MEADE, RN, PHD, AND JAMES C. BYRD, MD, MPH

Abstract: This study was designed to determine if primary care clinic patients read at a level congruent to the reading grade level of available smoking literature. Reading estimates of smoking education literature (N = 49) ranged from grade 3 to college level (median = 9.5). Reading skills of 258 smoking patients were estimated by reported years of schooling (median = 10), and by scores on the Wide-Range Achievement Test (median = 6). A serious disparity existed between the reading estimates of smoking education literature and the literacy skills of patients in our sample. (*Am J Public Health* 1989; 79:204–206.)

Introduction

Smoking is the cause of 350,000 premature deaths a year in the United States.[1,2] Smoking educational literature is commonly distributed to reinforce patient education and extend verbal instructions regarding smoking cessation.[3–7] Printed materials are suitable only when they can be read and comprehended.[8] There is strong evidence that a disparity exists between reading levels of written health materials and reading skills of target populations.[9–16]

Readability formulas are designed to make quick assessments of the readability of printed text.[17] Developers of written materials must also consider such elements as content, format, technical accuracy, organization, reader interest, and motivation.[8,18,19]

Educational level is often the only indicator used by practitioners to evaluate patient's comprehension skills. It has been suggested that reading skills may be two or more grade levels lower than stated educational levels.[10,11,14,15,20] Informal cues which may indicate that patients have low reading skills include blaming their inability to read on forgotten glasses or exhibiting facial signs of frustration while reading a passage.[21] More formal methods of estimating reading skills involve administering the CLOZE procedure[22,23] or the Wide-Range Achievement Test (WRAT).[24]

We compared reading levels of smoking educational materials with the reading skills of smokers.

Methods

Sources of written smoking educational materials (N = 49) were collected from organizations that supported smoking cessation activities: American Heart Association (N = 7), American Lung Association (N = 14), American Cancer Society (N = 7), government agencies (N = 12), and private

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organizations (N = 9). Three one-hundred word passages were randomly selected from each booklet. Estimates of the readability of each booklet were calculated by the Minnesota Interactive Readability Approximation Microcomputer Program (MNIRAP), a text analysis program intended to predict reading level estimates of educational materials. A computerized approach was selected over manual calculations of reading estimates of the smoking education booklets to increase the reliability of the results and decrease human error. The following formulas are calculated from the program: Dale-Chall, Raygor, Fry, Flesch, Gunning-FOG, and SMOG.[25]

The patient sample consisted of 258 subjects from the Primary Care Clinic of the Milwaukee County Medical Complex (a public hospital facility) who were participating in a smoking cessation program. All subjects were 18 years of age or older, English speaking, English reading, and had corrected vision that allowed for ease of reading. Each participant was asked to state their completed years of schooling and were then administered the WRAT Level 2. The WRAT consists of a graded list of words that yields a word-recognition grade level score. It is easily administered, scored in a manner of minutes, produces highly reliable and valid scores, and has been used in a variety of clinical settings.[10,11,14,24]

Results

The readability estimates of the 49 smoking education booklets ranged from grade 3 to scientific/professional level, with the majority of booklets written above grade 9. Table 1 illustrates the range of readability levels for a sample of five smoking booklets. Pearson Product Moment Correlations revealed that the formulas were positively and highly correlated with each other and ranged from (r = 0.75, 95 per cent CI: .59, .86 to r = 0.99, 95 per cent CI: .98, 1.0). Due to the strong correlations between formulas, one readability formula (SMOG) was selected for use as a comparison for the study results. This formula was selected because it is used widely in analyzing health literature and is a quick and easy method for estimating readability if a microcomputer program is unavailable.[10,14,18,26] According to the SMOG formula, there were 10 booklets written at grades 6 to 8.9, 27 booklets written at grades 9 to 11.9, and 12 booklets written at grade 12 and greater level. The mean reading level of the booklets according to the SMOG was grade 10.5 with 80 per cent of booklets written at grade 9 and greater (details of analyses available on request from author). Chi-square analyses revealed that there was no significant difference between the mean level of readability of the booklets obtained from the various public and private organizations.

The reported level of education of the subjects ranged from no schooling to college level with a median level of grade 10. Reading ability as measured by scores on the WRAT ranged from less than grade 3 to above grade 12 with a median level of grade 6. There was a moderately positive correlation between reported years of schooling and WRAT scores (r = 0.5550, 95 per cent CI: .47, .63). The mean difference between educational grade level and WRAT score was 3.8 grades which was statistically significant (t = 19.226, 95 per cent CI: 3.45, 4.21). Of the 90 subjects (35 per cent) who stated that they had completed high school, only 45 (18 per cent) demonstrated this level on objective testing. Although 77 per cent (N = 178) of the patients had a WRAT score of less than grade 9, only 23 per cent (N = 59) had stated educational levels less than grade 9. Fifty-four per cent of the subjects (N = 119) had reading levels (WRAT scores) less than grade 6.

A serious disparity existed between reading levels of patients and reading estimates of smoking literature (Figure 1).

Discussion

We have identified an incongruence between the reading level of written smoking educational materials and the reading skills of patients in our primary care clinic. Our results support the finding that the mean grade level of patient education booklets often ranges from grade nine to eleven, and are written at levels greater than the reading skills of the target population.[9,10,14–16,27,28] Results also indicated that agencies committed to smoking cessation activities write their publications at similar reading levels (grades 9–11).

Readability, predicted by a readability formula, is a necessary first step in the development of smoking educational materials. It offers an objective measure of text difficulty. Patient motivation, type, size, format, or use of graphics must also be considered.[29–31] Readability formulas do not take into account the fact that some longer words may increase clarity and be more familiar to the smoker e.g., cigarette and nicotine. Such factors must be considered when developing written texts for smoking patients.

Users of readability formulas need to consider that formulas have been validated in previous studies, but not on health care populations.[19,25,32–38] The scores obtained by the various methods for a given booklet may differ because of the measures used as predictors and the criterion by which the methods were validated. Future research needs to be directed to better identify if the SMOG is the most appropriate formula for estimating readability of health materials.

Years of schooling was not found to be a reliable predictor of reading ability. Other informal and formal indicators of reading skill must be used by practitioners to determine the patient's literacy skills. Incorporation of an objective test such as the WRAT is recommended for

TABLE 1—Readability Estimates of Smoking Educational Literature

Title	Author	Dale Chall	SMOG	Raygor	Fry	FOG	Flesch
The Decision Is Yours	American Cancer Society	13–15	12.3	College	11	12.9	12.2
Children and Smoking: A Message to Parents	American Heart Association	7–8	10.8	12	9	11	10.3
No More Butts. A Guide to Quitting Smoking	US Department of Health & Human Services	7–8	8.6	6	6	8	7.1
Chronic Bronchitis	American Lung Association	11–12	12.3	College	12	13	13
To Smoke or Not to Smoke	Channing I Bete Company	9–10	10	8	7	9.6	8.1

NOTE: Table 1 represents 10% (N = 5) sample of literature analyzed.

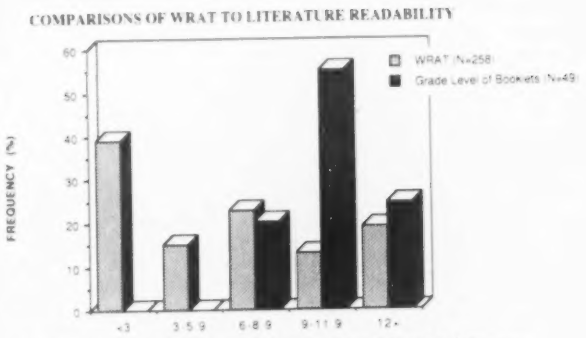

FIGURE 1—Comparisons of WRAT to Smoking Education Literature Readability

inclusion in a patient assessment. Patients in our setting did not appear threatened or frustrated by use of this instrument even when their reading skills approached illiteracy.

The median educational level of our patients was grade 10 which is lower than national statistics which identify grade 12.6 as the median educational level.[39] Individuals with less than a fifth grade reading ability (44 per cent of our sample) can be termed "illiterate",[40] and may not benefit from written materials. Other methods of smoking cessation education need to be explored and evaluated. These include use of flip charts, videotapes, audiotapes, picture books, or computer-assisted instruction.[20,41]

Evaluating both the readability of the smoking educational literature and the comprehension levels of patients allows practitioners to determine whether health information is compatible with the reading skills of their patients. This philosophy can help create positive smoking health messages that patients want to read, can read, and can understand.[8,19,26]

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Age of Initiation and Switching Patterns between Smokeless Tobacco and Cigarettes among College Students in the United States

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Abstract: A stratified, random, multi-stage, cluster sample of physical activity classes among 72 colleges and universities throughout the United States yielded 5,894 usable responses (males=2,888); (females=3,006). Twenty-three per cent of the students who began smoking before the age of 10 were current smokers, while 61 per cent of those who began using smokeless tobacco before the age of 10 were current smokeless tobacco users. College students were more likely to switch from smokeless tobacco to cigarettes than from smoking cigarettes to smokeless tobacco. (*Am J Public Health* 1989; 79:207–208.)

Introduction

A great deal of speculation has been focused on the age at which smokeless tobacco users initiate the habit and whether these young dippers and chewers eventually switch to cigarettes or whether the reverse pattern holds: i.e., cigarette smokers switching to smokeless tobacco. In oral testimony before the National Consensus Development Conference, Christen[1] and Hunter[2] testified that young people initiate smokeless tobacco and then switch to cigarette smoking due to the fact that, as users become older, smokeless tobacco becomes more socially unacceptable. Conversely, Glover[3] testified that due to the tremendous publicity surrounding the health hazards of smoking cigarettes, youthful smokers view smokeless tobacco as a safe alternative to cigarettes, and consequently choose to become dippers or chewers after initiating tobacco behavior with cigarettes.

The answers to these questions will require a longitudinal study. However, because the health consequences of tobacco (smoking or smokeless) use are profound, the following national collegiate cross-sectional study using retrospective questions was undertaken.

Methods

Students at colleges and universities in the United States enrolling 2,000 or more undergraduates were the focus of this study. The US was stratified geographically into eight regions, with one state from each region randomly selected to represent that region:

Central: Ohio
South Central: Oklahoma
North Central: Minnesota
Rocky Mountain: Colorado
Southwest: Arizona
Northwest: Oregon
Northeast: Connecticut
Southeast: South Carolina

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Thereafter, all colleges and universities in the eight states which met enrollment requirements of 2,000 or more students were contacted requesting their participation. Of the 109 schools contacted, 72 (66 per cent) chose to participate, five declined, two could not administer the questionnaire within the prescribed time, and 30 did not respond to repeated contacts.

In April 1986, at each participating college or university, random physical education activity classes were sampled; irrespective of their major, students were required to take activity classes in 67 of 72 colleges. The activity classes were randomly drawn from the most active classes (weight lifting) to the least active (bowling). The sample closely matched race and gender enrollments of each university population. The chairperson at each of the 72 participating schools was asked to survey approximately 100 subjects in a cluster of two to five classes. Several schools returned less than the 100 surveys requested, and 562 surveys were eliminated for failing to follow instructions or omitting portions of the survey.

The operational definitions for regular users were self-reports of current users using half a pouch (chewers), half a can (dippers), or half a pack (cigarettes) for more than three months. Age of initiation was defined as first age of regular use and not first experimentation.

The survey instrument, a 42-item questionnaire, contained a section soliciting demographic information followed by questions on cigarette smoking and smokeless tobacco use.[4] The data were checked and analyzed with the Statistical Analysis System (SAS), Version 5.

Results

The sample consisted of approximately equal percentages of male (49 per cent; n=2,888) and female (51 per cent; n=3,006) students. Mean age was 20.8 years. Eighty-six per cent (n=5,069) were White; there were 413 Blacks (7 per cent). Approximately one-third of the subjects were freshmen with decreasing numbers of sophomores, juniors, seniors, and graduate students. One-third of the subjects came from cities with populations greater than 50,000, approximately one-fourth from cities of 30,000-50,000, and the rest in decreasing order were from smaller towns, and rural (<5,000) areas.

The overall prevalence of cigarette smoking was 14 per cent and for smokeless tobacco products 12 per cent. Females were more likely to smoke than males (16 per cent vs 13 per cent) and males were more likely to use smokeless tobacco (22 per cent vs 2 per cent). The age of initiation of tobacco use among college students did not vary a great deal by tobacco type (cigarette vs smokeless products).

Data from questions regarding the age at which college students initiated cigarettes or smokeless tobacco and whether they currently used tobacco products revealed that those who initiated cigarette smoking at an early age were less likely to be current users than early age initiators of smokeless tobacco. In contrast, there was a strong association between age of initiating smokeless tobacco and current

TABLE 1—Age of Initiation of Tobacco Use vs Current Use of Tobacco by Gender and Tobacco Type

Age of Initiation (years)	Currently Smoking				Currently Using Smokeless			
	N	%Male	%Female	%Total	N	%Male	%Female	%Total
<10	91	31	14	23	56	60	63	61
10–12	99	42	54	46	78	74	38	71
13–15	247	72	70	71	184	93	50	91
>15	388	84	85	85	389	91	81	90
All	825	70	69	100	707	86	87	86

smoking. Of those who initiated either habit, most began at age 16 or older (Table 1).

While most current smokeless tobacco users do not smoke, those who began using smokeless tobacco prior to the age of 13 were proportionately more likely to be current smokers than those who did so at later ages (Table 2).

Most college students who smoked, did not use smokeless tobacco; however, the likelihood of smokers starting to use smokeless tobacco decreased substantially for those students who began smoking at ages 13-15 and 16 and older when compared with those who began smoking before the age of 10 or between the ages of 10 and 12 (Table 3).

Eighteen per cent of those who had started at any age had also used smokeless tobacco. Twenty-eight per cent of those who had ever used smokeless tobacco had also smoked cigarettes regularly. Among current smokers, 15 per cent occasionally used smokeless tobacco, and among current users of smokeless tobacco, 18 per cent occasionally smoked.

Discussion

In our survey the age of initiating tobacco use among college students did not vary a great deal by tobacco type (smokeless or smoking), contrary to two statewide studies of students grades 3 through 12 who started to use smokeless tobacco at earlier ages than cigarettes.[4,5] Presumably, the sudden upswing in popularity of smokeless tobacco over the past few years accounts for the difference. As the cohort of young initiators currently in elementary and secondary schools reach college level, the reported age of initiation may change. However, since educational level influences smoking, perhaps these two distinct populations (college vs secondary or elementary) will continue to show differing patterns of use regarding smokeless tobacco.

TABLE 2—Age of Initiation of Smoking and Use of Smokeless Tobacco among College Students Who Started Using Smokeless Tobacco

Age of Initiating Smokeless (years)	Age of Initiating Smoking								Do Not Smoke	
	<10 yrs		10–12 yrs		13–15 yrs		>15 yrs			
	%	N	%	N	%	N	%	N	%	N
<10	31	18	7	4	2	1	2	1	58	33*
10–12	6	5	22	17	8	6	8	6	56	44
13–15	3	5	3	6	6	11	9	17	79	145
>15	3	12	4	16	5	19	14	54	74	287

*Rows add to 100%

TABLE 3—Age of Initiation of Smoking and Use of Smokeless Tobacco among College Students Who Started Smoking

Age of Initiating Smoking (years)	Age of Initiating Smokeless								Do Not Use Smokeless	
	<10 yrs		10–12 yrs		13–15 yrs		>15 yrs			
	%	N	%	N	%	N	%	N	%	N
<10	15	14	4	4	4	4	9	8	68	62
10–12	3	3	14	14	4	4	12	12	67	66
13–15	0	1	2	5	4	10	6	15	88	215
>15	0	1	1	4	3	12	11	43	85	328

*Row adds to 100%

The smoking data on age of initiation of cigarettes and its relation to current smoking were also contrary to previous research which showed that smokers who initiated smoking earlier were more likely to continue smoking.[6] There is no ready explanation for this discrepancy. However, at the time of this data collection smokeless tobacco was very popular and this, as well as the nature of our sample, may have influenced our findings.

The 1986 Surgeon General's Report[7] and the 1986 Consensus Development Conference[8] were concerned that youngsters were turning to snuff and chewing tobacco as an alternative to cigarette smoking. While this may hold true for younger age groups, our data from college students suggest the switch was from smokeless tobacco to cigarettes. Moreover, most of the persons we surveyed will secure white collar jobs, where smokeless tobacco is not socially acceptable. No effort was made to validate self-reports; since we were interested in switching patterns, validating self-reports may have been less important.

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Public Health Then and Now

The Sociologic Context of Occupational Health in South Africa

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Abstract: The early history of the occupational health system in South Africa is outlined up to the early 1970s which mark a political and social transition in the society. Relevant demographic and social data are provided, and the roles and mutual relations of capital, labor, state, and academic sectors are discussed. During the past 15 years there has been heightened occupational health activity. Major legislative activity has included several commissions, the promulgation of new laws and regulations governing the workplace, and deregulatory measures in a contradictory mix. Conflictual relations between social forces are illustrated by two examples involving the introduction of safety representatives in the workplace, and compensation for occupational lung disease. The implications of wider political and economic realities are analyzed, and current and probable future trends in the evolution of the occupational health system are identified. (*Am J Public Health* 1989; 79:213–221).

Introduction

This paper seeks to present an account of the occupational health situation in South Africa. Part I is an historical documentation of the evolution of the occupational health system, covering both political and legislative aspects up until the early 1970s, which mark a political and social transition in the society; Part II consists of a demographic and sociologic profile of the society in relation to occupational health; Part III deals with the relatively underdeveloped occupational health infrastructure in South Africa.

Part IV is the core of the article. Since the 1970s there has been heightened occupational health activity, and it is the social dynamics of this most recent period that are of concern. Legislative activity has included several commissions and the promulgation of new laws and regulations governing the workplace. Relations between management, labor, and the state have expressed both inequalities of power and conflicts of interest with regard to the generation and implementation of new legislation. Wider political and economic realities have also had serious implications for health and safety. Recent events are reviewed with the purpose of identifying current and probable future trends in the evolution of occupational health activities.

I: An Historical Overview

Sociopolitical Factors

The history of South Africa is steeped in racism. This has affected all aspects of the body politic, and underlies the development of the occupational health system.

Industrial development began at the end of the 19th century with the discovery of diamonds and gold. The mining industry brought with it new patterns of industrial and political relations in which White mine owners, White skilled mine workers, and both Black and White unskilled mine workers were thrown together in a situation fraught with potential conflict and compounded by language-group hostilities among the Whites. Craft-based unions organized by White workers excluded Blacks on grounds of skill and race.

Persistent attempts by management to eliminate mining skills and to replace White with cheaper Black labor together with poor working and health conditions on the mines[1] led to a period of sustained conflict in the first two decades of the 20th century, culminating in an unsuccessful insurrection by White miners in 1922. The ruling party which supported the interests of mine owners was replaced shortly thereafter by a coalition of agrarian-based Afrikaner nationalist and White labor parties. Black South Africans were consequently excluded not only from the political structures of government but also from legal status as workers and from access to trade unions and industrial bargaining structures. This was the mechanism whereby the monopoly rights of Whites over skilled trades in both mining and non-mining industry were entrenched.[2]

A new wave of mainly Black trade unionism occurred during World War II when industry was highly dependent on Black labor due to the development of secondary industry stimulated by the war economy. This period was associated with a significant updating of occupational health and compensation legislation. However, in the 1950s and 1960s, Black labor demands were suppressed along with the wider political demands of Black South Africans by the enforcement of apartheid under the nationalist government that came to power in 1948.

The development of occupational health in South Africa has always been prompted by labor activity. Organized White labor succeeded in obtaining many concessions from capital such as preventive legislative and compensation legislation

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for miners' phthisis and work-related accidents. Compensation benefits for pneumoconiosis among White mine workers were in advance of conditions enjoyed by workers in the developed countries of Europe and North America as far back as 1956 (The Pneumoconiosis Act, 57 of 1956). After separate but parallel struggles over health issues, Black workers succeeded in obtaining some coverage, although benefits were invariably racially discriminatory.

The role of production workers of different color changed profoundly over time. Most White workers took on supervisory and managerial roles which distanced them socially from Black production workers as well as from exposure to production hazards. White workers' racism and the incorporation of White labor unions into segregated state industrial relations structures led to the emergence of White-dominated or White-exclusivist trade union federations with a bureaucratic character.[3-5] These trade unions have often obstructed Black labor organization; they have tended to side with management or the state when direct political repression has been directed at democratic non-racial unions. To some extent, because of their experience with these organizations, Black workers have tended to develop a negative view of trade unions.

Legislation

South African occupational health legislation has generally followed two parallel tracks, one covering the mining industry, and the other dealing with non-mining industry, commerce and services.

Various laws relating to mining health and safety preceded the formation of the Union of South Africa in 1910. In 1901 the Government Mining Engineer in the Transvaal reported high mortality rates from miners' phthisis (a combination of tuberculosis and silicosis). This was followed by the first Commission of Inquiry into Phthisis in 1902, and three years later by the first mining regulations created to abate the dust hazard. The regulations were not particularly effective, and another commission, the Mining Regulations Commission of 1907, was set up to investigate dust control. The Commission's report in 1910 showed that the mortality of White miners was six times higher than that for adult males on the Witwatersrand. At this time small *ex gratia* payments were made by the mining companies to widows of victims of miners' phthisis.

The Colonial laws that existed prior to Union were consolidated into the Mines and Works Act, 12 of 1911, which was intended to provide for preventive measures and the protection of the health and safety of mine workers. The Act related mainly to the control of machinery. A further commission of medical practitioners was established to look into miners' phthisis and tuberculosis and to make recommendations for compensation. This resulted in the Miners' Phthisis Act of 1911 which introduced compulsory compensation for phthisis. These two Acts led, after various commissions and amendments, to the two main Acts currently in operation—the Mines and Works Act, 27 of 1956, and the Occupational Diseases in Mines and Works Act, 78 of 1973. These Acts provide for the control of the work environment on the mines, for risk or fitness examinations relating to miners' fitness for underground work, and for benefit examinations for occupational disease compensation.

Outside of the mining industry, the Workmen's Compensation Act No 25 of 1914 and an amendment of 1917 provided the first coverage for industrial accidents and occupational diseases, respectively. Disease had to be present along with disablement, and had to be causally related to work. Posthumous benefits were available only if death was caused by the occupational disease. Cyanide rash, lead and mercury poisoning were the three occupational diseases recognized and they were handled administratively as if they were accidents. These were augmented by ankylostomiasis in 1934, and by 1941, 15 occupational diseases including silicosis were scheduled as occupational diseases.

The first Factories Act was passed in 1918, and from 1924 industry was administered by the Department of Labor. Prior to this the mines and industry had been jointly administered by the Department of Mines and Industry. Limited coverage for white collar workers in offices was provided by the Shops and Offices Act, 41 of 1939. The new Factories, Machinery and Building Work Act, 22 of 1941, replaced the outdated 1918 law; it laid down basic conditions such as hours of work and regulations pertaining to the control of machinery. For the first time a Factories Inspectorate was constituted within the Department of Labor, with duties to inspect workplaces and investigate reportable accidents.

In 1958, the Minister of Labor appointed a departmental committee to investigate and make recommendations on the incidence of occupational diseases. The committee reported back in 1963 providing evidence of widespread occupational disease in South African industry, but the report did not result in any substantial preventive or compensation measures.

The Shops and Offices Amendment Act, 75 of 1964, introduced new health and safety-related protective measures relating to hours of work and other conditions of employment for white collar workers in shops and offices. The Factories Act was amended again in 1967 and some general health regulations were introduced but, with the exception of regulations relating to noise control, no regulations specifically dealing with occupational hazards were promulgated until the mid 1980s.

II: Demographic and Social Profile

Population

South Africa has a population of over 30 million which is divided by law into the following racial categories: Black Africans, people of mixed racial descent, Asians, and Whites (Table 1). The African segment is by far the largest, comprising about 80 per cent of the total population.[6] The social class breakdown demonstrates that the Black African population consists mainly of unskilled and semi-skilled workers, and the White population mainly of white collar and managerial personnel, while the Asian and mixed-race populations fall somewhere in between—closer to the Black African group than the White group[7] (Table 2).

A high proportion of the African workforce are migrant workers who commute back and forth from rural dormitory areas in the interior. These rural reserves were created by the 1917 Land Act which allocated a small proportion of South Africa's land for the majority African population. In recent years some of the reserves have been designated "independent" ethnic or tribal states by the South African government in terms of its Bantustan policy. These areas have no economic viability apart from their relationship to South Africa. No African peasantry remains after a century of migrant labor, dependence on earnings in the wage economy of White South Africa, overcrowding of the reserve areas, generalized impoverishment, and high unemployment.

TABLE 1—Demographic Profile of South Africa: Population Estimates 1986 (in millions)

Demographics	Population in millions (%)
South Africa (including the independent homelands)	
Total Population	32.0
Total Black Population	26.0
Average Black Economically Active	12.5
Average Black Unemployed	6.5
Unionized Blacks	
Paid up	1.1
Signed up	1.5
The Independent Homelands (Transkei, Bophytatswana, Venda, Ciskei)	
Total Population	6.8
Economically Active Population	3.7
Unemployed	1.8
Industrial Sector Distribution (1980 Population Census)	
Manufacturing	1.5 (21)
Mining	0.8 (11)
Agriculture	1.3 (18)
Construction	0.4 (5)
Commerce	1.3 (18)
Services	2.0 (27)
Total Employed	7.3 (100)
Total Economically Active	10.0

Union Federations Membership	1986–87	Signed up	Paid up
COSATU	1.86 1.87	0.6 1.0	0.4 0.7
NACTU	1.2.86	0.4	0.3
TUCSA	1.2.86 87	0.1 0.0	0.1 0.0

SOURCE: Data from Sarakinsky M, Keenan J (reference 6).

Capital

South Africa is a relatively well-developed capitalist country. There is an important locally based large capital sector, in addition to the presence of foreign capital. These sectors tend to favor independent labor organization and more representative political dispensations. There are also significant medium and small capital sectors, and a large state production sector which is conservative in its labor policies. More recently there has been a concerted attempt to encourage the development of an informal sector (family production and trading) and a small capital sector; this attempt takes the ideological and political form of a campaign to deregulate the economy.

Labor

By third world standards, South Africa has a large working class, comprising about one-third of the total population. After a 10-year lull, the Black working class manifested its presence again in the early 1970s through a series of strikes involving large numbers of workers in important industrial centers of the country. This heralded the beginning of today's independent trade union movement.[8]

Trade unions—independent of the management and racially based structures characteristic of the state industrial relations system prior to the 1980s—now have a national presence in most industries. The two national federations are the Congress of South African Trade Unions (COSATU) and the National Council of Trade Unions (NACTU). Estimated membership figures are given in Table 1. There are approximately one million members of the independent trade union movement,[9,10] about 10 per cent of an economically active population of 10,000,000. Although this represents only a modest proportion of the total workforce, membership has increased rapidly in recent years, and is relatively strong in some of the most crucial sectors such as mining, metals, and chemicals (Table 1). The old White-dominated Trade Union Council of South Africa disbanded at the end of 1986 after experiencing rapidly declining membership in the 1980s due to major splits and desertions by component trade unions with large numbers of Black members. As Table 2 indicates, although there are trade unions for the small number of residual "White workers," this social category hardly exists.

Because the independent workers' movement has grown from below in open conflict with state, capital and race-segregated or White-dominated trade unions, there have been opportunities for grassroots or rank and file organization to develop at shop floor level, and the possibility of workers being incorporated into unrepresentative joint structures with representatives of the state and management (such as joint health and safety committees) has been minimal. By contrast with other African countries whose national trade union bodies have appeared within a year or two of independence in situations of little rank and file organization in the factories, it took 12 years of organization to establish COSATU in 1985 by which time factory organization was fairly solidly based.[11] Despite the severe economic effects of the recession, disinvestment, massive retrenchments of workers in most industries, and political repression since 1985, much of which has been aimed at the independent trade unions and COSATU, there has been an exponential increase in membership.[9] Health and safety issues have been successfully negotiated with managements.[12–14]

Other Black Social Groupings

Although there is no longer a significant self-supporting peasantry in the rural areas, a dormitory working class exists; resettled urban unemployed, ill, and old workers and women and children live in rural "towns" with no access to the means of subsistence production. There is also a significant

TABLE 2—Percentage Composition of the South African Population by Race and Social Class, 1969–77

Social Class	United Kindom	South Africa			
		White	Mixed Race	African	Total
Upper	5	11	0.6	0.5	4
Middle	30	60	25	13	30
Working	65	29	75	87	68
Skilled	38	23	13	5	11
Semi-skilled	18	5	29	20	17
Unskilled	9	1	33	62	40

SOURCE: Data from Hindson D, Simkins C (reference 7).

FIGURE 1—Women in the Ciskei (homeland) carrying water for their village.

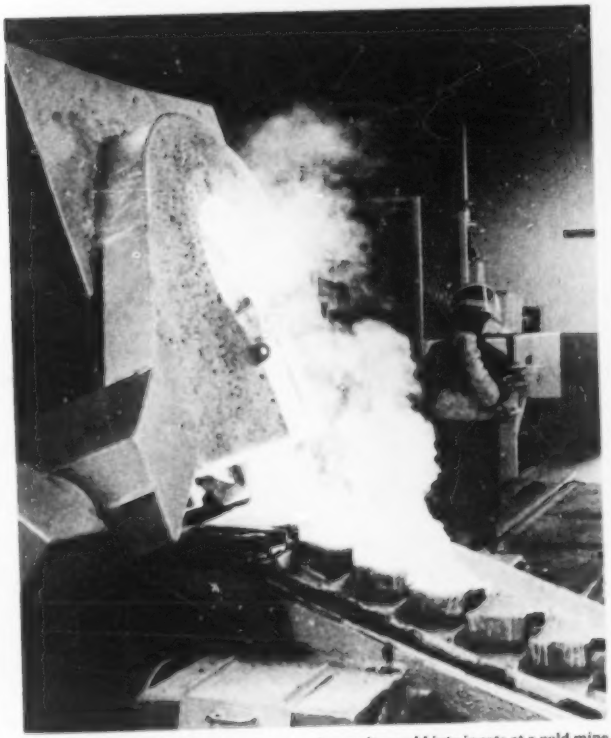

FIGURE 2—South African worker pouring molten gold into ingots at a gold mine in Welkom, Orange Free State.

and fast-growing middle class consisting of white collar workers, professionals, government employees and functionaries, small-business operators, and small traders. This process of class transformation is being actively assisted by the state and capital, especially large capital.

The State

Since the implementation of the new constitution in 1984, White, mixed-race, and Indian South Africans have been separately enfranchised in racially exclusive representative bodies. The majority of workers, Black Africans, are not enfranchised. As Black workers generally perceive the state to be hostile to their interests, it has been very difficult for the state to incorporate their organizations. Lack of political representation has in fact been an important stimulus for the Black working class to develop alternative channels of political representation through their trade unions by winning respect from management, putting forward their ideas and demands, and by gaining rights and conditions of work exceeding those provided for in the law.

Before 1979 the state followed a course of suppression of Black labor organization. The 1979 Wiehahn Commission elaborated a more sophisticated and systematic attempt by the state to control the independent trade union movement and to extend "normal" industrial relations structures in the hope that this would lead to the incorporation and containment of Black trade unions and labor militancy.[15,16] During the period 1979 to 1984 there were protracted conflicts between organized labor and the state[17] around such issues as the registration of trade unions, state control and incorporation into the state industrial council system. The state lost ground in these conflicts: restrictive race-segregation conditions had to be dropped from registration criteria for unions, and independent unions have generally been able to join industrial councils on their own terms and utilize their status and presence to the advantage of their members, while reserving their right of non-participation.

It may be argued that a combination of antagonism between the state and Black workers and the availability of occupational health benefits won historically by white labor have allowed the independent trade union movement to take advantage of such concessions without being easily subject to domination or incorporation.

III: The Occupational Health Infrastructure

The State Sector

The Medical Bureau for Occupational Diseases in Johannesburg operates a clinic where it executes its statutory duties under the Occupational Diseases in Mines and Works Act, 78 of 1973, relating to fitness and benefit certification examinations for White miners. Some epidemiological research is also conducted, and this has increased in recent years. The Bureau certifies Black miners with occupational diseases on the basis of their medical records and findings that are forwarded to it by local doctors on the mines, but Blacks are not examined on the premises. The Bureau is also the main medical certifying authority recognized by the Workmen's Compensation Commissioner for cases of occupational disease falling under the Workmen's Compensation Act, 30 of 1941.

The National Center for Occupational Health in Johannesburg runs a clinic where workers from non-mining industry who are suspected of having occupational diseases may be evaluated. This clinic does not have statutory functions like its counterpart at the Medical Bureau, and access to the clinic is not racially restricted. A combined panel of doctors from the Center and the Bureau evaluate cases seen at the

Center's clinic for purposes of certification; these certifications are also recognized by the Workmen's Compensation Commissioner. The Center is involved in basic and applied research into occupational medicine, industrial hygiene, pathology and toxicology, and has a well-equipped library which provides a national information service.

The nationwide National Occupational Safety Association is financed out of the Accident Fund established under the Workmen's Compensation Act. It is involved in voluntary safety audits of industrial firms and in providing information and educational activities related, until very recently, only to occupational accidents. Since the passage of the Machinery and Occupational Safety Act there has been a move to expand its educational activities to encompass occupational disease prevention.

The Corporate Sector

A substantial network of health services, consisting of hospital and general practitioner facilities, is associated with the mining houses. From an occupational health perspective, these services are geared mainly to emergency general health problems and occupational accidents and their sequelae.

The non-mining industry has a very poor record of occupational health services. A survey of occupational health services in the Greater Cape Town area found a low percentage (12.9 per cent) of workplaces with any health service at all (the lowest level of service was a part-time nurse).[18] A similar percentage was found in Germiston, which is one of the most heavily industrialized areas in the country.[19]

The Academic Sector

Occupational health is a neglected area within the medical schools. There is little time devoted to it in the undergraduate curriculum and as yet it is not a registered specialty. Departments of Community Health at several universities, principally the Afrikaans-language universities, have been active in the area for many years, and part-time postgraduate diplomas in occupational health are available at the Universities of Witwatersrand (in conjunction with the National Center for Occupational Health), Cape Town, Stellenbosch, and Pretoria. At Stellenbosch the medical school runs a rehabilitation clinic, and at Cape Town the medical school has a respiratory clinic for occupational diseases.

The Departments of Sociology at the Universities of Cape Town and Durban have research groups active in occupational health research, education, and consultancy. These are the Industrial Health Research Group and the Industrial Health Unit, respectively.

There is no university level industrial hygiene or environmental engineering training in the country. Industrial hygiene is currently part of the health sciences curriculum at technical colleges in major urban centers where diplomas in health inspection are offered.

Independent Service Organizations

The Work Information Group in Johannesburg and the Industrial Health and Safety Education Project in East London are active in occupational health research, information, and education.

The Trade Union Sector

Several of the independent trade unions have been active in health and safety. This activity has taken the form of union-run clinics (Food and Allied Workers Union and the Southern Natal branch of COSATU; health and safety departments and health and safety officers (National Union of Mineworkers); union medical officers (Food and Mineworkers unions); health and safety organizational structures such as shop stewards committees or safety representatives, health and safety educational activities for union safety personnel, and negotiated health and safety agreements conferring rights of access to health and hazard data. Unions are increasingly involved in direct and commissioned industrial hygiene and occupational health research activities and provide input to the government Department of Manpower on proposed health and safety legislation and regulation. Labor has drawn upon the resources of the various sociology departments and other independent organizations in many of these activities.

Occupational Epidemiologic Research

There is little reliable information about the extent of occupational disease in South African industry. Nationwide statistics are hard to obtain and the quality of available mortality and morbidity data for African workers cannot support meaningful epidemiologic studies.[20] Some statistics of variable quality for the mines are published by the Medical Bureau. The Workmen's Compensation Commissioner does not publish any data on occupational disease. Data are available for occupational accidents and show high mortality and morbidity, especially on the mines. The profiles of the older occupational diseases like pneumoconioses and metal poisoning have been documented by government sources as being high.[21] Newer occupational diseases like cancer, neurological damage, and musculo-skeletal conditions arising from chemical, solvent, and radiation exposures and poor workplace design are as yet undocumented and unstudied.

Historically occupational epidemiologic activity in South Africa has been linked to health problems associated with mining and, with a few notable exceptions, has related to dust-induced occupational diseases. Studies, not often published, have mostly originated in the statutory bodies (the Bureau and the Center). The bulk of published material has been in the form of case reports, pathology studies, and collections of statutory data with few prevalence or incidence studies. They have also been based mainly on data for White workers who constitute a very small and relatively unexposed section of the workforce especially in the mining industry.

The output of research from universities and corporations has been low, despite the existence of collections of routine data, particularly in the large firms.

Since the end of the 1970s, published output from South African occupational health research has increased somewhat. Much of this is collaborative research with independent trade unions. Research activity has advanced from clinical or pathological post-mortem case reports or series and cursory screenings to more sophisticated and carefully designed epidemiologic studies.

IV: Recent Events

The Commissions of Enquiry

Labor turbulence during the early 1970s led to a flurry of legislative activity and three state commissions of enquiry were established to investigate labor-related legislation: *The Erasmus Commission of Enquiry on Occupational Health* reported in 1975 on its investigations into the occurrence of occupational diseases, statutory protection of health at work, and occupational health services.[22] The findings were that:

- occupational diseases were widespread and highly prevalent in many industries;
- statutory protection of health at work was fragmented, incomplete, and out of date; and
- occupational health services outside of the mining industry were very limited.

Many specific recommendations were made with the major implications being that all workers should be covered by protective legislation and that a uniform industrial health policy and regulatory practice should be obtained. It was proposed that overlapping functions and costs should be eliminated by amalgamating all control within the Department of Health under the provision of one Act; that workers and their representatives be consulted about working conditions affecting their health; that employers be obliged to provide industrial health services, safe working environments and health information; and that an occupational disease database be established. The Commissioners stressed that in their view the situation was parlous and urgent and that speedy action was required. They emphasized that it was important to rectify poor working conditions before matters were taken up as organizational issues by Black labor.

The Wiehahn Commission of Enquiry into Labor Legislation reported in 1979.[15] The Commission was requested to bring the system of labor relations and the dispute-handling machinery into line with changed circumstances, and was charged with producing methods for constructing sound labor relations in South Africa. Workplace health and safety was a major area of enquiry. The main recommendations were that

- there should be a participatory dispensation for all parties and races in the economy;
- consultation and cooperation should take place between employers, trade unions and employees before changes were made or new developments introduced in the workplace; and
- industrial relations should be deracialized.

It was advocated that participation in the compensation process should take place by means of a Workmen's Compensation Board consisting of employer and employee representatives, which would review the provisions of the Act and its administration. The Commission's opinion was also that mining industry legislation should be deracialized with respect to both prevention and compensation.

The Nieuwenhuizen Commission of Inquiry into Compensation for Occupational Diseases reported in 1981.[23] The objectives of this Commission were to investigate and make recommendations relating to the differences between the two Acts in respect of all aspects of ascertainment, diagnosis, determination of disability caused by occupational diseases and the financing of compensation; and to look into the question of rationalizing legislation and practices. The Commission recommended integration of the two systems under the format operative in the mining sector. It advocated the tightening-up of compensation criteria in anticipation of the deracialization of the proposed new legislation and increased numbers of Black claims.

New Acts

New Acts and Bills in the area of occupational health and safety were not introduced until 1983: the Basic Conditions of Employment Act, 3 of 1983, the Machinery and Occupational Safety Act, 6 of 1983, and the Occupational Medicine Bill.

The Machinery and Occupational Safety Act was inspired by the legislation of other countries, especially the British Health and Safety at Work Act (1974) and attempted to modernize legislation relating to the protection of health and safety. As a piece of enabling legislation, it held out the possibility of moving toward increasingly stricter regulation of safety at work and the adoption of various standards relating to the use of hazardous substances (see below).

In many respects, the two new Acts only went part of the way toward meeting the recommendations of the Commissions. While legislation was modernized and deracialized, the objective of having a single Act was not achieved. Sexual discrimination was removed from the legislation by the Basic Conditions of Employment Act, but so was the protection (special health provisions) of women workers. None of the new laws included the trade unions as participants in the reforms and in the new structures that were introduced. The organizational structures laid down in the Machinery and Occupational Safety Act for factory health and safety were not consistent with the spirit of the Commissions as far as consultation with employees and trade unions was concerned.[24]

The new legislation signals the emergence of a changing perspective embodying contradictory trends. In some respects worker health and safety rights have been weakened, while in other respects new rights have been conferred and added responsibilities placed on employers.[25,26] Concern with regulating conditions must also be seen in conjunction with recent moves toward deregulation which is dealt with below. The removal of race and gender criteria from the body of labor legislation (including health and safety) has not constituted a significant reform since the composition of the labor force has in any case changed over the years resulting in a workforce that is today mainly Black with substantial female participation.

The Occupational Medicine Bill of 1983 proposed the establishment of factory-based health services along the line of the Erasmus Commission proposals. Unfortunately the Bill was inadequately drafted; it would have duplicated the factory inspectorate set up under the Machinery and Occupational Safety Act, and was ultimately dropped.[27]

During 1986, a Bill was put forward proposing amendments that would remove job reservation from the notorious Mines and Works Act, one of the most explicitly racially discriminating pieces of labor legislation on the statute books. Changes to the racial allocation of jobs could have an important effect upon labor participation in health and safety structures and practices on the mines. The National Union of Mineworkers, which has campaigned since its inception in 1983 for deracializing laws and practices on the mines, gave evidence to the Standing Parliamentary Select Committee on Mineral and Energy Affairs on their objections to the Bill. One of their main reasons for opposing the Bill was the wide discretionary power granted to the Minister, which would enable the retention of racial discrimination by ministerial proclamation despite the formal removal of racial criteria for holding skilled mining jobs.

Officials at the office of the Compensation Commissioner for Occupational Diseases in terms of the Occupational Diseases in Mines and Works Act have intimated* that attempts are currently under way to remove the racial basis for compensation awards under this Act, and to make

*Personal communication from the Office of the Commissioner.

earnings the basis of compensation. New measures flowing from the Nieuwenhuizen Commission are anticipated in the near future.[28]

While there have been some insignificant amendments to the Workmen's Compensation Act since 1981,[29,30] important compensation practices have been administratively altered behind the scenes, disadvantaging claimants (see below).

New Regulations

To date, a small number of regulations under the Machinery and Occupational Safety Act have been drafted. Some of these regulations are new, others are updated versions of pre-existing regulations. Only the General Safety, General Administration of the Act, and the Asbestos Regulations have so far become law. Regulations governing environmental conditions (noise, heat, and cold), general machinery, driven machinery, workplace facilities, and electrical machinery have been pending for long periods now. The regulations have generally taken a long time to pass from draft to final form and there has been concerted management resistance to new control measures, resulting in a weakening of successive drafts.[31] In some cases, notably the Facilities Regulations, the draft indicates a potential substantial weakening of benefits to workers when compared with regulations dating back to the 1960s or the 1940s. However, in the case of the general safety regulations, industrial workers are given some important new rights, and additional obligations are placed on employers. Thirteen years on the proposals for regulations in the Erasmus Commission report have by no means been adequately met.[32]

Deregulation

The Temporary Removal of Restrictions on Economic Activity Act was enacted in 1986, and represents a disturbing development with potentially severe implications for health and safety, as well as for wages and other conditions of work. In terms of this Act, laws may be suspended by the State President in any geographic area or industrial sector. In the field of health and safety for example, specified employers could be exempted from all legal responsibilities. Some limited protection of health and safety regulations is afforded by the stricture that the President must act on the advice of a standing committee of Parliament and in consultation with the Minister of Manpower. However, protection would only be exercised effectively by union lobbying at this level.

The underlying purpose of this Act appears to be the promotion of free enterprise and small business, in particular by removing "restrictive legislation" by presidential decree. It is very likely that this Act will be applied to the most vulnerable, unorganized workers, whose conditions are generally the worst.

The implementation of deregulation via legislative initiatives has been accompanied by a great interest in privatization, which appears to be the dominant trend in state and management economic thinking. The contradictory picture that has emerged of worsening conditions and practices, alongside new regulations and deregulation seems to reflect conflicting thinking in the state with regard to an economic goal of privatization, and the objective of an acceptable industrial relations environment.

Social Forces in Conflict: Two Examples

Both management and the state have been alarmed at working class activities since the early 1970s. Labor-linked health and safety activities have increased steadily and there have been a series of management and state responses calculated to preempt worker organization, disrupt production, and erode managerial prerogative at the workplace. These conflicts will be illustrated by means of two examples: the safety representatives and safety committees of the Machinery and Occupational Safety Act, and struggles over worker compensation for occupational lung disease.

Example 1: The Act, Safety Representatives and Safety Committees

The response of state and management to the Act is an instructive lesson in the sociology of shop-floor primary health care. The law is imprecisely worded and thus permits different views regarding its implementation. According to the Act safety representatives are to be designated in writing by management and this has been interpreted by management to mean that they are management appointees; it is also assumed that safety representatives should be literate (most Black workers are not) and should have completed training courses certified by the National Occupational Safety Association before being eligible. There also has been a strategy of appointing only supervisors and managerial personnel, who are usually White, as safety representatives.[24] This has resulted in production workers and their elected representatives being bypassed.

Organized labor on the other hand has endeavored to democratize the application of this law by negotiating the right to elect safety representatives, by reserving the right to negotiate health and safety issues through representative shop stewards' committees, by avoiding the management-dominated joint safety committees, and by extending rights in health and safety agreements far beyond the provisions of the Act.

Example 2: Compensation for Occupational Lung Diseases

Since the end of 1984, changes in Workmen's Compensation practices have taken place which place workers at a disadvantage. In the last few years large numbers of compensable occupational disease cases, particularly pneumoconiosis, have been generated in the wake of union-related health screening programs.[33–36] The reaction of the compensation authorities has been to make benefits for occupational disease less accessible and less favorable to workers. In 1984 the Workmen's Compensation Commissioner prevented the joint Bureau-Center certifying panel from certifying cases as compensable disease for submission, unless cases were seen personally at the Center's clinic in Johannesburg.[37] This meant that it was no longer possible to have the medical particulars of suspected cases outside of the Johannesburg area forwarded for certification to the joint panel prior to submission to the Commissioner for compensation.

In late 1985, the disability classification for workers with occupational chest diseases was revised administratively by the Commissioner. Before that time, workers with radiographic evidence of pneumoconiosis as their only manifestation of an occupational disease were assessed as 35 per cent disabled; this was the lowest category of disablement for occupational lung diseases and entitled workers in this category to a lifelong pension. After the revision, the lowest disability category was set at 20 per cent and entitled workers in this category to a lump sum payment only. In money terms, the value of the benefit was thereby reduced to between a third and a half of its former value depending on the age of the worker.[38–40] This disadvantageous revision was introduced with no reference to organized labor and attempts to raise the matter directly with the Commissioner or indirectly via Parliament proved unsuccessful.

The revision was immediately applied by the state medical certification bodies. Observations from industries like the asbestos-cement industry are that virtually all cases of asbestosis now certified are receiving the lump sum payment instead of the pension. A further disadvantage is that the company pensions which had previously supplemented the Commissioner's pension to 100 per cent of the last basic wage, and allowed workers with asbestosis to opt for voluntary retirement on these terms, are being withheld from those now classified as only 20 per cent disabled. Total compensation in these cases consists of the lump sum from the Commissioner.

Conclusion and Future Implications

The past 15 years have been characterized by a process of tentative governmental political and labor reforms in South Africa. However, political exclusion of Blacks, privileged relations between management and state, and the underlying hostility of the state towards labor which has frequently erupted despite legislative reform, have created a negative climate for co-optation of Black labor militancy. The state has failed to fob off workers' political demands with substitute trade union rights.[41]

In the field of occupational safety and health there has been a process of reform characterized by a technical updating of legislation appropriate to new developments in industrial production. The social content of this reform is conservative: labor has been excluded from a participatory role at the various levels of regulation. Workers and their organizations remain suspicious of reform initiatives that do not specifically include them in a representative way.

A number of new legislative measures evince contradictory state thinking. On the one hand there are new regulations which rely on management self-regulation and are in some cases weaker than pre-existing legislation, and on the other hand there are deregulatory measures. Changes to compensation practices have been put into effect, affecting workers with chronic lung disease in dust-producing industries; similar changes are embodied in draft legislation currently being considered by parliament. Early indications[24] that the Machinery and Occupational Safety Act enshrined the principle of management self-regulation and the exclusion of labor have been borne out by the method of its implementation and the method proposed for monitoring new draft regulations. There is little evidence of any increase in the activity or efficacy of the factory inspectorate. For the foreseeable future, the state appears to be headed for a minimalist role in occupational health, with increased legislative activity but only formal and weak control over the work environment.

The new legislation will most probably lead to an increase in the numbers of occupational health care professionals. Indications are that they will be inserted into a privatized network of services paid for and directly controlled by management.

The role of labor has been mainly confined to taking up issues at a shop floor level. At the national level, some progress has been made in negotiating health and safety conditions in the asbestos industry, while the National Union of Mineworkers has made important submissions to the parliamentary standing committee charged with redrafting the Mines and Works Amendment Bill. Neither the COSATU nor NACTU have thus far developed a health and safety policy or program, although one is currently being considered by COSATU. The past three years have been associated with weakened labor health and safety activity as a result of the depletion of union administrative resources by state action against union members and officials, including two States of Emergency, as well as the effects of the economic recession. The absence of worker participation in the legislative process could lead to a situation where legislation and practices now being established may be seen as technical and socially neutral, not requiring transformation in a possible post-apartheid South Africa, and survive into the future relatively intact.

There appear to be compelling reasons for coordinated labor responses to halt the creeping disaster of deregulation, and to influence the modernizing of health and safety regulation which is now taking place. However, one of the effects of the States of Emergency has been to harden anti-state feeling on the part of workers and their organizations, and this could reinforce the tendency to avoid dealings with the state as a matter of principle. The important question is whether unions will follow the example of the mineworkers and engage the state through the national labor federations with the aim of influencing the legislative process. This may be appropriate, particularly in areas such as changes to the compensation procedure, or attempts at deregulating certain sectors of the economy, where more direct and participatory forms of worker action are very difficult. In these instances, direct labor input on key measures at the national federation level may be the only way to decrease labor's vulnerability to the designs of management and the state.

Management has been able in many instances to use the recession and the disinvestment and sanctions campaigns to place unions on the defensive in wage and other bargaining by portraying labor and other popular organizations as the root causes of disinvestment and sanctions. One example in the health and safety area is that of halting the practice of substituting dangerous substances like asbestos in construction materials on the grounds that imported materials are more costly. Poor exchange rates, which are inextricably bound up with the political situation, were blamed. In another case, the management of a chemical factory where a number of workers had been poisoned claimed that they could not afford effective protective measures as their foreign parent company had decided to disinvest.

In most workplaces management has made changes during the past two years to comply with the legal requirements to implement health and safety structures laid down by the Act. However, working conditions in many factories still contravene legal requirements, and the topdown approach to the implementation of safety structures produces a strong impression that the Act has been extensively implemented in form rather than in depth.* Delays in the passing of draft regulations suggest management opposition to major changes aimed at improving health and safety, and concerted efforts to promote their interests in the regulation process. This would be consistent with the overall developing economic program, shared by the government and capital, comprising conservative measures such as privatization, deregulation and other anti-labor economic policies, which simultaneously represent a threat to safe and healthy working conditions. These responses reveal the way in which health and safety is continually weighed up against other cost factors to the

*Macun I: The Implementation of the Machinery and Occupational Safety Act. Paper No 8 XVIIIth Annual Congress of the Association for Sociology in Southern Africa, Cape Town, 1987.

enterprise, rather than being viewed as an important worker right.

Occupational health is a significant aspect of the well-being of a population. This paper has attempted to show the importance of the participation of worker-patients in the amelioration of occupational health problems historically and currently, and the linkage between labor organization and improvements in the work environment. The conservative nature of occupational health system reform in South Africa and the current trajectory of labor suppression threaten to exclude labor as a participant in this system. This raises the question as to whether labor's occupational health gains will be maintained or lost in the future.

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The American Association of Public Health Dentistry (AAPHD) has issued a call for abstracts for its 52nd annual meeting, to be held November 1–3, 1989 in Honolulu, Hawaii. Papers on a broad range of dental public health subjects are invited, and a special call was sounded for international reports, due to the setting and the theme.

The theme of this years AAPHD meeting is "Commitment and Caring: The Challenge for the Next Century." Key topics suggested by the program planners include: trends in oral disease; innovative preventive and educational measures; national and local dental programs; dental care delivery systems; ethics and access to care; oral health for special population groups; effects of technology on dental public health; AIDS; infection control; and other related topics. *Abstracts are due by April 1, 1989.*

For further information, and rules for submitting an abstract, contact the AAPHD Program Chair, E. Joseph Alderman, DDS, MPH, Director of the Office of Dental Health, 878 Peachtree Street, NE, Room 215, Atlanta, GA 30309.

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both federal and local levels between public health workers and others with legitimate concerns in this field.

This issue of the Journal features a wide range of papers addressing both intentional and unintentional injuries as well as basic research and evaluation of intervention efforts. Such a gathering of papers is possible because more public health workers have been entering the field, in many cases through new or revitalized formal programs in health departments and schools of public health. Their efforts are beginning to bear fruit. It is noteworthy that this issue also contains a research contribution from France, reflecting in part recent efforts by the World Health Organization to promote injury control research and public health programs throughout the world.

It is axiomatic in the field of movie-making that the sequels are never as good as the original. In the field of public health, however, we have substantial hopes—and some promise—that "Injury Control IV" will be the best yet.

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ERRATA

In: Hatziandreu EI, Koplan JP, Weinstein MC, Casperson CJ, Warner KE: A cost-effectiveness analysis of exercise as a health promotion activity. Am J Public Health 1988; 78(11):1417–1421.

On page 1419, last paragraph, right column, line 7, the sentence should read: Coronary artery bypass graft surgery with left main disease costs $5,000 per QALY, but this figure increases to $40,000 when only two vessel disease and mild angina are present.

* * * * * * *

In: Goodman SN, Royall R: Evidence and scientific research. (Commentary) Am J Public Health 1988; 78(12):1568–1574.

On page 1574 in the Appendix, there is a typographical error in the second equation in relation to the positioning of the square root symbol. The square root should be positioned only over the "N" as follows:

The Z-score is:

[2] $Z = 4/\sigma_p = 4\sqrt{N}/20 = \sqrt{N}/5$.

In the published version in the December Journal, the square root was erroneously extended over the 20 and the 5. We apologize for the error.

Survival Experience of Aged Hip Fracture Patients

JAY MAGAZINER, PHD, MSHYG, ELEANOR M. SIMONSICK, PHD, T. MICHAEL KASHNER, PHD, J. RICHARD HEBEL, PHD, AND JOHN E. KENZORA, MD

Abstract: Hip fracture has long been considered a major threat to survival in aged populations. This report describes the survival experience of 814 aged, community dwelling hip fracture patients treated in seven Baltimore hospitals between 1984 and 1986: 4.3 per cent died during hospitalization; 8.2, 12.6, and 17.4 per cent died within three, six, and 12 months after fracture, respectively. The mortality rate for the entire population approaches expected mortality approximately six months post-fracture, but varies by age and sex. The most important factors predicting mortality are presence of serious concomitant illness and marked delirium (in the absence of dementia) at the time of hospital admission. The authors suggest that medical factors that may contribute to patient disorientation be investigated and treated, when possible, in an effort to improve the survival status of hip fracture patients. (*Am J Public Health* 1989; 79:274–278.)

Introduction

This report describes the survival experience of a recent series of hip fracture patients who entered one of seven Baltimore area hospitals between October 1, 1984 and April 30, 1986. While the survival experience of numerous series of hip fracture patients has been previously reported, the majority of these include patients treated in the late 1950s to mid-1970s,[1–25] generally before the initiation of prophylactic anticoagulant and antibiotic treatment for surgical patients.[13] Secondly, most recent hip fracture series involve patients treated outside the United States,[26–34] where treatment principles and use of technology differ markedly from those in this country.[35] Moreover, most previous reports do not differentiate the mortality experience of individuals residing in nursing homes or other long-term care institutions at the time of fracture from that of community dwelling patients,[1,3,5–11,13,15–17,19,20,22,24,25,31] and numerous studies[1,2,5,6,8,9,11,13,17,19–23,31,32,36,37] include hip fracture cases younger than age 55 for whom the cause of fracture is often severe trauma and for whom the prognosis of recovery is generally excellent.[3,24,25,37]

The present study provides more current information on the survival experience of aged hip fracture patients treated in the United States, and residing in the community at the time of their fracture, and determines at three, six, and 12 months following hip fracture the relative risk of mortality associated with selected demographic and medical indicators.

Methods

Subjects and Data Collection

Subjects consist of all patients aged 65 years or older, admitted to one of seven Baltimore area hospitals from the community with an acute fracture of the hip between October 1, 1984 and April 30, 1986. We identified the majority of patients (n = 698) through weekly calls to an established contact within each hospital, usually an orthopedic ward nurse or an admissions officer. To ensure complete enumeration of eligible cases, we reviewed hospital discharge records for cases discharged with an ICD-9 code of 820 not previously identified. An additional 116 cases were located in this manner. Data on all patients (n = 814) were obtained from hospital charts by trained interviewers using a standard protocol.

The seven study hospitals were selected on the basis of the volume of patients treated using data from the Maryland Health Services Cost Review Commission. These hospitals treat approximately half of all hip fracture patients age 65 and older treated within the 27 hospitals in the Baltimore SMSA. Study hospitals comprise a broad range of practice settings, including urban, suburban, sectarian, and public hospitals. Three of the seven hospitals have orthopedic residency programs. Elderly hip fracture patients treated in the study hospitals compared with those seen in all Baltimore area hospitals are more likely to be older than 75 years (79 per cent versus 75 per cent), female (83 per cent versus 79 per cent) and discharged to nursing homes (39 per cent versus 34 per cent).

Measures

Predictors—The demographic factors examined in this study include: age, sex, and race. The medical predictors include: type of fracture (intracapsular or extracapsular), orientation on admission, and concomitant medical conditions. The measure of orientation on admission derives from physician chart notes regarding patient status on admission. If disorientation was noted, the patient was labeled confused; if unnoted, the patient was considered mentally clear.

To account for patients' medical status pre-fracture, we constructed a binary measure of general medical condition based on whether any of the following conditions were noted in the patient's chart: cancer; stroke and related conditions; heart, atherosclerotic, liver, bladder, pulmonary, or renal disease; chronic obstructive pulmonary disease (COPD); or Parkinson's disease. These conditions were selected on the basis of previous findings involving single disease entities and mortality following hip fracture[1,9,16] and on clinician judgment. We constructed a binary measure of cognitive status based on whether or not senile dementia, Alzheimer's disease, or organic brain syndrome were noted in the hospital chart. Using this measure and an additional item concerning orientation on admission, we developed a three-category variable. The reference category includes patients with neither dementia nor noted delirium on admission; the second, consists of only patients with dementia; the third, includes cases with noted disorientation on admission, but who were not diagnosed as suffering from an organic brain syndrome. We chose to distinguish between chronic confusion due to dementia and acute confusional states or delirium

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as the latter often indicates the presence of a potentially serious biomedical condition or metabolic disturbance which can deleteriously affect patient outcome.[17,38]

Outcome—We obtained the survival status of patients and date of death (when applicable) through follow-up interviews of the patient or a close friend or relative of the patient at two, six, and 12 months following hospital discharge. For 119 patients for whom we could not obtain follow-up interviews and the 116 patients identified through discharge diagnoses, the investigators searched the death records maintained by the Maryland Bureau of Vital Statistics for each patient's name and date of birth. When a match was found, staff recorded the date of death. If no death record was found, the patient was presumed still living.

Using the date of death for patients who died and the date of hospital admission, we calculated the number of days each patient survived since sustaining a hip fracture. Since hospitalization generally occurs within one or two days following a hip fracture,[22] we used the date of hospital admission as a reasonable approximation of the date of fracture.

Data Analysis

We calculated the expected one-year mortality using 10-year age group-, sex-, and race-specific death rates published by the National Center for Health Statistics for the United States population in 1982.[39] We calculated the observed mortality rate for each month following fracture and compared this with the expected death rates. In plotting the expected survival curve and in calculating the expected monthly mortality, we assumed that change in mortality is directly proportional to time over the one year follow-up period.

To examine the relative predictive value of the demographic and medical factors, we first tested for the presence of an association between each variable of interest and mortality at three, six, and 12 months post-fracture, using Yates' chi square statistic for dichotomous measures and Mantel-Haenszel's test statistic for variables with three or more ordered categories. We performed logistic regression analyses with mortality as the outcome for each time period after fracture. Odds-ratios and 0.95 confidence intervals were calculated for each factor using the beta coefficients and standard errors generated by the regression procedure following the method prescribed by Schlesselman.[40]

Results

Table 1 describes the demographic and medical characteristics of 814 patients followed for one year after sustaining hip fracture. The population consisted largely of White females with an average age of 80 years. Patients were relatively unhealthy with over 70 per cent having one or more serious concomitant illnesses and 8.6 per cent suffering from organic brain dysfunction. Thirty-seven patients (4.3 per cent) died before hospital discharge—all but one within 40 days of admission.

The proportion of patients dying within three, six, and 12 months following hip fracture is 8.2, 12.6, and 17.4 per cent, respectively. Figure 1 plots the observed and expected one year survival curves for the total population and for the population subgroups. Hip fracture patients clearly experience greater mortality than persons of similar age, race, and sex. The degree of excess mortality, however, varies by sex and age. At one year the difference between the observed and expected mortality rates is greater for males than females and greater in patients younger than 75 years than older patients. The point at which the observed survival curve parallels the expected survival curve also varies by age and by sex. The death rates for female patients and patients 85 years or older approach expected rates approximately six months post-fracture. Patients between 75 and 84 years of age experience similar mortality as expected at about 10 months after sustaining hip fracture. Elevated mortality for both males and the youngest patients appears to continue beyond the first year post-fracture.

TABLE 1—Demographic and Medical Characteristics of 814 Hip Fracture Patients, Ages 65 and Older Admitted to Seven Baltimore Area Hospitals, October 1984 to April 1986

Characteristics	Per Cent (Years)
Gender	
Male	20.2
Female	79.8
Race	
White	93.5
non-White	6.5
Age	
Mean	(80.1)
65–74	24.2
75–84	45.3
85 or above	30.6
Type of Fracture	
intracapsular	45.2
extracapsular	54.8
Concomitant severe physical illness	70.6
Dementia noted in medical chart	11.8
Delirium on admission	19.3
Delirium on admission, exclusive of dementia	7.5

Table 2 presents the relative risk of mortality at three, six, and 12 months following hip fracture associated with the seven demographic and medical factors examined. The presence of at least one life-threatening condition as noted in the patient's medical chart confers the greatest risk of mortality relative to the other factors examined at three and six months following fracture. This elevated risk diminishes with increasing time since fracture from 4.6 to 2.6 times the risk experienced by patients wtih no concomitant illness. Patients with dementia do not experience an increased risk of mortality following fracture; confidence intervals of any apparent protective effect of dementia are very wide, however. Patients with noted delirium on admission who had no history of dementia have 3.1 to 3.5 times the risk of dying as cases who were not disoriented when admitted to the hospital or who had a history of cognitive impairment.

Patients older than 84 years of age have 2.6 times the risk of dying within three months of hip fracture as patients between ages 65 and 74 years, decreasing by six months after fracture. Patients between ages 75 and 84 years do not experience an increased risk of mortality relative to patients between ages 65 and 74 years. These findings, with respect to age, are particularly noteworthy given that the expected death rate for persons 85 years of age and older is nearly six times the mortality rate for persons between ages 65 and 74 years (14.6 per cent versus 2.5 per cent) and the expected death rate in persons 75 to 84 years old is more than twice that of the rate for persons between ages 65 and 74 years (5.8 per cent versus 2.5 per cent). As expected the relative risk of mortality associated with male sex increases from 1.4 to 1.9 over time.[39] The small proportion of non-White cases (6.5 per cent) makes any finding, regarding race, difficult to interpret.

The relative risk associated with each factor after adjusting for the other factors in the models does not differ greatly from the unadjusted relative risk, except for a slight reduction in magnitude evident for each factor across all time periods.*

*Data available on request to authors.

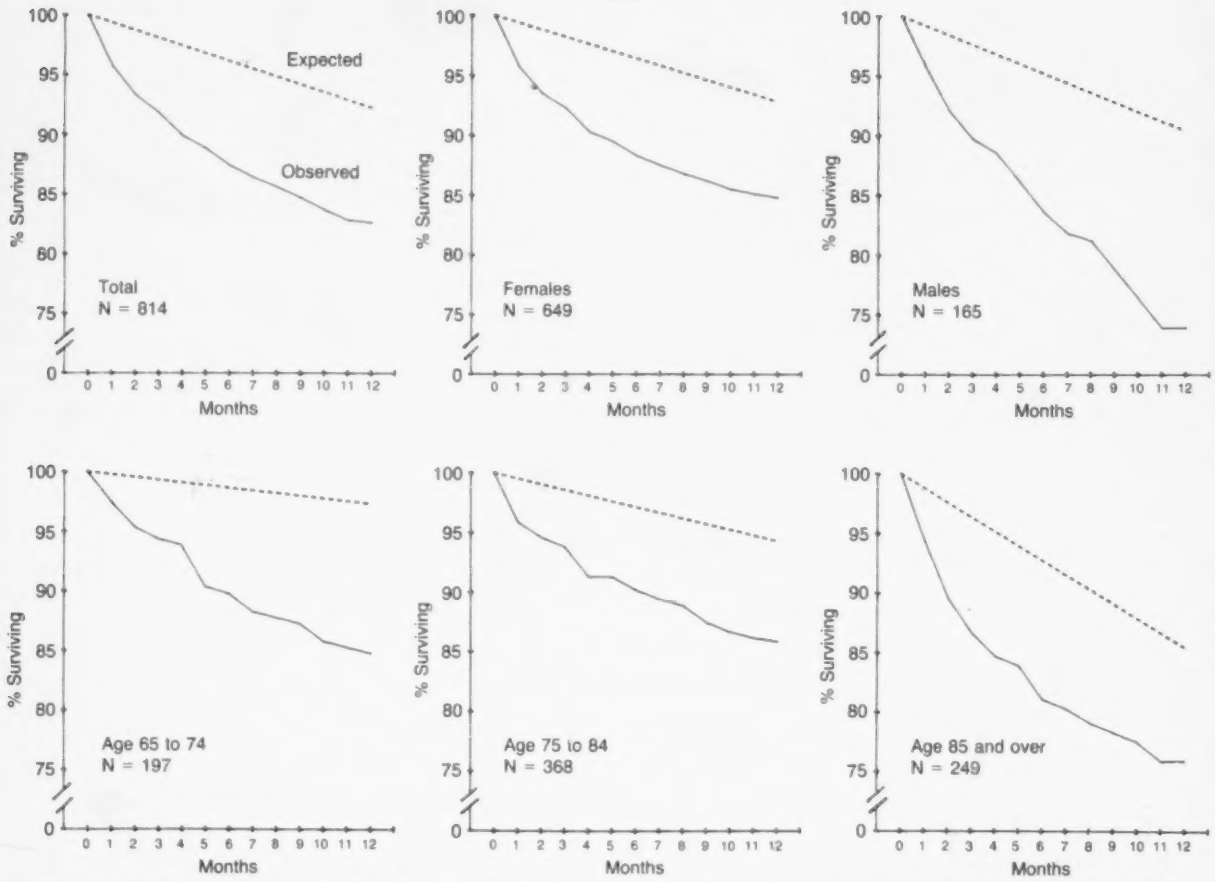

FIGURE 1—Observed and Expected Survival Curves by Sex and by Age Group for 814 Hip Fracture Patients, Ages 65 and Older, Baltimore 1984 to 1986.

Figure 2 presents the probability of one year survival given patient sex, age category, illness state, and orientation at the time of hospital admission independent of dementia. Delirium is present in a relatively small proportion of patients—7.5 per cent overall and no greater than 16.2 per cent in any subgroup. Patients older than 84 years and patients with at least one serious medical condition are more likely to become notably disoriented following their fracture than younger and healthier patients. In comparing subgroups similar in all respects except orientation on admission, we find that in all instances where valid comparisons can be made, patients with delirium on admission experience a lower probability of survival than patients with either normal orientation or a history of dementia. For female patients with no serious concomitant disease, the observed survival rate does not differ from the expected rate in either young or old patients. This relation does not hold for male hip fracture cases. Lastly, patients with serious medical problems have a lower probability of survival than patients without any serious medical conditions whether or not they become disoriented.

Discussion

Studies of mortality following hip fracture are characterized by diversity in patient populations, length of follow-up, and country of treatment. This diversity renders assessment of recent secular trends in post-fracture survival difficult, although it appears that mortality rates among hip fracture patients have generally declined between the late 1950s and the mid-1970s.[13] As such, the comparison of findings with previous work is limited to case series treated in the 1970s and later.

The in-hospital mortality rate of the current series of fracture patients is 4.3 per cent, considerably lower than the 18 and 21 per

TABLE 2—Unadjusted Relative Risk of Mortality at Three Points in Time Following Hip Fracture Associated with Selected Factors in 814 Patients Ages 65 and Older, Admitted to Seven Baltimore Area Hospitals, October 1984 to April 1986

Factors	Months Since Hip Fracture		
	Three	Six	Twelve
Sex (male)	1.4 (0.8,2.4)*	1.5 (0.9,2.4)	1.9 (1.3,2.9)
Race (non-White)	1.5 (0.6,3.6)	1.9 (0.9,3.8)	1.8 (0.9,3.3)
Extracapsular fracture	1.2 (0.7,2.0)	1.2 (0.8,1.8)	1.1 (0.7,1.6)
Ages 75 to 84**	1.1 (0.5,2.4)	1.0 (0.5,1.7)	0.9 (0.6,1.5)
Ages 85 and older**	2.6 (1.3,5.3)	2.1 (1.2,3.6)	1.8 (1.2,2.9)
Concomitant disease	4.6 (1.9,10.8)	3.6 (1.9,6.6)	2.6 (1.6,4.2)
Noted dementia	0.4 (0.1,1.7)	0.8 (0.4,1.7)	1.5 (0.9,2.5)
Delirium (exclusive of dementia)	3.2 (1.6,6.3)	3.5 (1.9,6.3)	3.1 (1.8,7.5)

*95% confidence interval in parentheses.
**relative to patients ages 65 to 74

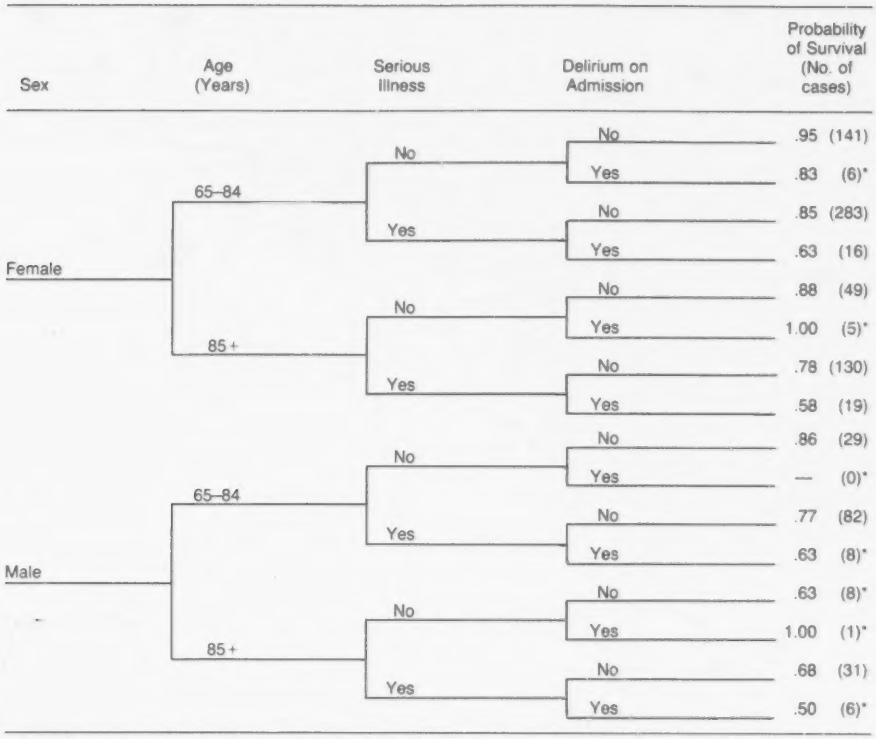

FIGURE 2—Probability of Survival One Year Following Hip Fracture Given Certain Characteristics
***Interpret probabilities with caution as these subgroups have fewer than 10 cases.**

cent mortality found for two similarly aged patient series treated, respectively, in Great Britain[26] and in Ireland[27] in 1981. The length of hospital stay for those patients averaged over 30 days, however, compared with less than 20 days for the current series. Our in-hospital mortality rate (4.3 per cent), compares favorably to the 2.0[32] and 5.8[36] found in two other recent series treated in Scandinavia that included patients as young as 50 years of age. Differences in in-hospital mortality across studies may reflect gross differences in treatment procedures, health status of patients, and/or age distribution of patients treated. While none of the case series described above include institutionalized patients, criteria for placement in and the definition of long-term care facility may vary greatly across cultures.

The one year mortality of patients in the current series (17.4 per cent) is similar to that found in two series, treated in Denmark between 1977 and 1978, that exclude non-elderly patients. One[36] had a six-month mortality of 11.1 per cent; the other,[28] 20 per cent. Differences in observed mortality reported in other studies[29,32,37] appear strongly related to the age distribution of the respective patient populations reviewed.

The observed and expected survival curves for the current series become approximately parallel six months post-fracture, but this point varies by age and sex. This finding generally agrees with earlier studies which report that a substantial portion of excess mortality occurs in the first few months following fracture.[35] Estimates from other studies of patients hospitalized since the late 1960s[5,24,28] of the point at which the mortality curve begins to parallel that for a cohort matched by age and sex range from three to 12 months.

Several investigators[1,15,19] have found a secondary rise in mortality at varying points following hip fracture ranging from four to six weeks to four to five months. Inspection of the survival curves and monthly mortality rates by age group in the current series suggests a similar phenomenon. This may reflect random variation in mortality rate rather than a true change in slope, but the possibility that a true secondary rise in mortality exists warrants exploration.

We found male sex, advanced age, presence of at least one serious illness, and delirium at the time of admission (exclusive of dementia) to represent threats to survival in the first year following hip fracture. Several previous studies report similar findings, with regard to sex, age, and concomitant disease.[1,3–5,13,24,28,33] Contrary to our finding that dementia is not associated with a higher rate of mortality, other investigations[1,24] find the presence of organic brain syndrome highly predictive of death within one year of hip fracture. These earlier reports, however, include patients from extended care facilities who may have had more advanced cases of dementia than the present community dwelling population; this difference could account for their higher death rates. While no previous reports examine the relationship between delirium at the time of hospital admission independent of dementia, the results from two prior studies are suggestive. Matheny and associates[17] found that patients who became disoriented during their hospitalization experienced much higher in-hospital mortality than patients who remained mentally clear throughout their stay (20.7 per cent versus 1.1 per cent). Barnes, *et al*,[4] examined blood urea (whose elevated levels are associated with acute confusion)[41] found the death rate among patients with elevated levels (>80 mg per 100 ml) died at five times the rate of patients with low levels (<40 mg per 100 ml).

How hip fracture contributes to delirium and why this

acute confusional state increases risk of mortality are important questions. While this area remains largely unresearched with respect to hip fracture cases, two reports offer some insight. Matheny, *et al*,[17] contend that post-fracture and post-surgical delirium may signify blood sugar or electrolyte aberrations, renal failure, infection, drug intoxication, among other clinically relevant states. Kenzora and associates[13] further suggest that patients who experience a delay in receiving medical attention may become dehydrated and that surgery before restoration of blood, fluid, and electrolyte balance may overwhelm the patient's capacity to make adequate physiologic adjustments. Taken together, these observations imply that mortality associated with patient disorientation may be amenable to treatment. Our findings tend to support this premise as marked disorientation on admission was found to make an independent contribution to patient mortality.

Several caveats are in order:

- The first concerns the dependence on medical chart notes for the measurement of dementia, confusion, and other significant medical conditions. The possible problems include: underreporting of serious medical problems, inability to evaluate the severity of various disease states, and mistaking delirium at the time of admission for dementia and visa versa.
- Second, the study population may not represent the universe of aged hip fracture patients residing in the community, as study hospitals were selected on the basis of the volume of patients treated and not for their case mix.
- Similarly, the exclusion of nursing home patients necessarily excludes patients in the poorest health. Thus, conclusions based on the current study population may not apply to all hip fracture cases.
- Fourth, the inability to determine the contribution of various disease conditions to mortality in the general population renders it impossible to distinguish the true contribution of hip fracture to patient mortality over that expected for a person with a similar health profile.
- Lastly, patient delirium noted on admission may have preceded the fracture—not followed from it.

Despite these limitations, our data indicate that excess mortality associated with hip fracture depends primarily on pre-fracture disease state and secondly on entering the hospital in an acute confusional state. Clearly, patients' pre-fracture medical status cannot be modified; however, fracture-related delirium is potentially treatable. This finding argues for careful investigation of factors that contribute to delirium as part of the standard treatment strategy.

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Determinants of Recovery 12 Months after Hip Fracture: The Importance of Psychosocial Factors

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Abstract: The independent contributions to recovery from hip fracture of psychosocial factors including depression, personality, social connectedness, and self-rated health were studied in 219 women age 59 and older (mean age 78.5) who were community dwelling prior to fracture. Initial assessments were conducted shortly after surgery and follow up assessments 2, 6, and 12 months later. By 12 months, 15 patients had died and 15 had entered a nursing home. Substantial declines in physical functioning though not psychosocial status were observed. Only 21 per cent (compared to 81 per cent prefracture) reported walking independently; fewer than 30 per cent had regained reported prefracture levels of physical function. The proportion with elevated depression scores at 12 months was 20 per cent, down from 51 per cent following surgery; 64 per cent rated their health excellent or good at 12 months, up from 43 per cent after surgery.

Poor cognitive status and post-surgical self-rated health were predictive of mortality. Among survivors, age, prefracture physical functioning, and cognitive status were associated with recovery in physical function but not psychosocial status. High post-surgery depression scores, but not the other psychosocial factors, were associated with poorer recovery in both functional and psychosocial status. These findings demonstrate the importance of depressive symptoms as one determinant of recovery from hip fracture and support the need to attend to the affective status of hip fracture patients following surgery. (*Am J Public Health* 1989; 79:279–286.)

Introduction

Fracture of the hip represents a public health problem of increasing magnitude. Annual incidence in the United States is estimated to be approximately 98 fractures per 100,000 population.[1–3] The majority of hip fractures (70–80 per cent) occur in females over 60 years of age, with the population at highest risk being White females 85 years or older.[4] While there are sophisticated surgical procedures for hip fracture repair, full recovery has been elusive. Only 12 to 23 per cent return to prefracture ambulatory status and/or functional independence.[5,6] Mortality rates in the first 12 months following fracture are 12 to 20 per cent higher than in comparable age/sex groups in the general population,[3,7–9] and rates of institutional placement during the immediate post-hospital period are high and remain higher than average as the individual grows older.[10,11]

Case series data[6,9–11] consistently have shown prefracture physical and cognitive function and age to account for a substantial amount of the variance in hip fracture recovery. While there is evidence suggesting that psychological, social, and environmental factors influence recovery as well,[6,12] evaluation of the contribution of a broad range of factors has been rare. Information regarding both physical functioning and psychosocial status, measured either before or after fracture, usually has not been available. Moreover, most studies have included both seriously debilitated, older individuals with limited potential for any recovery, and "healthier" elderly with a more positive prognosis. Consequently, even when psychosocial variables have been studied, there has been limited ability to evaluate the independent importance of these factors for recovery over and above the effects of the individual's health status.

We undertook a study designed to specifically investigate the role of psychosocial factors measured after surgery in the recovery of the "healthier" hip fracture patient. In this paper, we characterize their recovery status 12 months following surgery. We also present the findings of analyses to evaluate the independent contribution to recovery of four psychosocial factors measured during the initial, post-surgery assessments: self-rated health, depressive symptomatology, personality, and social supports. There is considerable evidence supporting the prognostic significance of these factors for the clinical course and/or recovery associated with diverse illnesses. The reader is referred to the excellent review by Kasl[13] for a critical appraisal of this literature.

Methods

Study Sample

Subjects were recruited from among community dwelling, ambulatory White females over 59 years of age who were treated surgically for a fractured hip in one of 17 hospitals in the greater Philadelphia area between August 1, 1984 and January 15, 1986. Initial structured interviews were completed as soon after surgery as possible (usually the 7th-20th day). Follow-up interviews were obtained two and six months later and a final telephone interview was conducted 12 months after surgery. Clinical assessments of walking ability and neuromuscular status two and six months postsurgery were also made, and the patient's hospital record was systematically reviewed.

Seven hundred and thirty individuals meeting the above age, sex, race, and residence criteria were identified at the time of surgery; they represent an almost complete ascertainment of the patients treated during the surveillance period in each hospital. Additional study criteria included: ability prior to fracture to walk across a room with no greater assistance than a straight cane, absence of post-surgical cognitive impairment that would preclude completion of study assessments and absence of cancer or other health problems likely to cause a pathological fracture or to result in death during the year following hip fracture. Three hundred and sixty-two patients met these criteria and were asked to participate in the study; 219 of these consented.

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The sample for this analysis, consists of the 211 individuals (96 per cent of those enrolled in the study) for whom complete post-surgical and follow-up data or documentation of death were available. Mean age of these individuals was 78.5 (s.d. 7.6). Prior to hip fracture, 55 per cent lived alone, 21 per cent lived with their spouse, and 24 per cent with friends or other relatives. Forty-eight per cent were of British or Irish ethnic origin; 25 per cent reported German descent, and 15 per cent and 11 per cent, respectively, were from Eastern Europe or were listed as other. Mean level of education was 11.6 years (s.d. = 3.4). Forty-three per cent of the study subjects sustained a fracture of the trochanteric or intertrochanteric region of the femur. An intercapsular fracture was reported for 36 per cent and a fracture of the subtrochanteric or femoral neck region was recorded for the remaining 21 per cent. For 65 per cent of subjects, hip fracture repair involved fracture reduction and fixation; for 35 per cent a partial or complete prosthesis was used.

Some information obtained during study enrollment was available for both the 211 patients in this analysis and those who refused the initial interview (n = 143) or who were excluded because of missing data (n = 8). In addition, survival status at 12 months was known for 94 per cent of the 143 refusers and 103 (72 per cent) completed the 12-month interview. Comparisons of study participants and non-participants revealed no differences ($p < 0.05$) between the groups in terms of: age, prefracture living arrangement, prefracture independence in ambulation, history of recent fracture, prefracture pain or observed physical or sensory impairments. At 12 months, there were no differences between participants and non-participants in terms of survival or any other recovery indice.

Study Variables

In addition to information on sociodemographic variables such as age, marital status, and education, variables pertinent to this report include:

- *Prefracture Health Status:* The number of health problems experienced in the preceding year was assessed during the initial post-surgical interview using a modified version of the health problems scale in the Multi Level Assessment Instrument (MAI).[14] Two additional questions were included to assess the presence and constancy of pain prior to hip fracture. Information abstracted from the patient's hospital record included the number of hospitalizations in the two years preceding the hip fracture and the number and type of different diagnoses reported on the admission history. Data on the type of diagnoses were used to classify individuals according to the number of serious diagnoses reported.
- *Prefracture Functional Status:* Ability to perform self-care activities of daily living (ADL) such as bathing, and dressing, and instrumental activities of daily living (IADL) such as shopping and doing housework prior to hip fracture were assessed retrospectively during the initial post-surgical interview. An adaptation of the Multi-level Assessment Instrument was used.[14] ADL and IADL scores were computed as the number of activities that the individual reported doing without any assistance. Difficulty in performing activities such as pushing, pulling, and carrying objects was measured with three items adapted from disability scales developed by Nagi.[15] The presence and kind of assistance required in walking or traveling distances were also determined.

  Predictably, the above measures of prefracture functional status were correlated. To minimize potential problems resulting from multicollinearity and to simplify the analytic models, a principal components analysis that included the function scales was done. This yielded a single factor. Factor loadings were used to compute a summary prefracture physical function score. Scores ranged from −4.33 (poor function) to .92 (good function).
- *Cognitive Function:* The Short Portable Mental Status Questionnaire (SPMSQ)[16] was administered during the initial interview to assess post-fracture cognitive function in terms of orientation, memory and abstraction ability. Of the 10 SPMSQ items, eight were asked; the questions on year of birth and age were omitted.
- *Psychosocial Variables:* Two subjective measures of global health status based on responses to the question, "How do you rate your health: excellent, good, fair or poor?", were obtained during the initial post-surgical interview. For one, individuals were asked to rate their health as they remembered it to be "just prior" to their hip fracture; for the other, a rating was obtained for current health status "today". The presence of depressive symptomatology experienced the week preceding the post-surgery interview was assessed with the Center for Epidemiological Studies Depression Scale (CES-D).[17,18] Recent evaluation of this 20-item scale confirms its validity as a measure of depressive symptomatology with older individuals, even those who are physically ill or disabled.[19] Scores range from 0 to 60 with a score of 16 or higher distinguishing those with a high probability of clinical depression.

  Personality dimensions of neuroticism and extraversion were measured by the Eysenck Personality Inventory.[20] This instrument contains 52 "yes/no" questions of which 24 are markers for each personality trait. The Eysenck Inventory has been extensively used in personality research. Recent reviews confirm the validity of the neuroticism scale and, to a lesser extent, the extraversion scale in older populations.[21,22]

  Social support, defined in terms of the availability and usual frequency of contact with children, close relatives, close friends and a "special person" with whom intimacies are shared was assessed using questions adapted from the National Institute on Aging project "The Establishment of Populations for Epidemiologic Studies of the Elderly" (EPESE).[23]
- *Treatment-Related Variables:* Fracture site (e.g., trochanteric, subcapital, intertrochanteric), repair type (pin or prosthesis), pre- or post-surgical medical events (e.g., fever, wound infection, pneumonia, congestive heart failure, decreased hemoglobin), number of days in hospital and discharge location (e.g. rehabilitation hospital, nursing home, home) were obtained from the patient's hospital record.
- *Recovery Parameters:* Recovery at 12 months was characterized in terms of survival status, physical functioning, self-rated health, presence of pain in the involved limb, extent of depressive symptomatology and return to prefracture level of physical functioning. Scales in the initial post-surgery interview to assess prefracture walking ability, ADL, IADL and physical

performance were readministered at 12 months with slight modifications to accommodate the telephone interview format. Again a principal component analysis was done to summarize the physical function scales. As before, a single factor emerged, and the associated factor weights were used in computing the 12-month physical function factor score. The global self-health rating referenced to "today", the pain questions and the CES-D were also readministered at 12 months. To accommodate the telephone format, a dichotomous response option, (never = 0; sometimes to always = 1) was used with the 20 CES-D items.

An additional recovery variable was constructed as a count of the number of seven different functional parameters on which the individual had returned to, or surpassed, prefracture levels. The criteria variables included return to prefracture: 1) residence location, 2) use of ambulatory aids, 3) ambulation in the home, 4) ambulation outside the home, 5) performance of self-care, 6) instrumental activities of daily living, and 7) participation in social and recreational activities.

Results

• *Prefracture Health and Functioning:* Although "healthier" hip fracture patients were enrolled, the study sample was not free of pre-existing medical conditions. Patients reported 3.5 (s.d. 2.2) health problems with a greater number of pre-existing medical conditions (x = 6.3, s.d. 3.1) recorded in the hospital record. Prevalent conditions noted in the hospital record included, coronary heart disease (60 per cent), hypertension (53 per cent), osteoporosis (51 per cent), and arthritis (46 per cent).

Most patients, however, reported independence in physical function prior to fracture. Only 19 per cent of the subjects reported requiring any assistance with ambulation. Most sample members (89 per cent) reported they were completely independent in ADL activities and 65 per cent indicated they completed all IADL activities without assistance.

• *Initial Psychosocial Status:* As seen in Table 1, a large proportion of study subjects (57.6 per cent) rated their health on the interview day as fair or poor. This is substantially in excess of that reported in large samples of community dwelling elders.[24,25] The retrospective ratings of health status (just prior to fracture) were much more likely to be positive. Over 50 per cent of the study sample scored 16 or above on the CES-D. Compared to most samples of community dwelling elderly, this proportion is extremely high. For example Berkman, *et al*,[19] report 15 per cent of women 65–74 years of age and 18 per cent of women ages 75+ enrolled in the EPESE study to have CES-D scores of 16 or greater. It is even in excess of the 28 per cent prevalence reported by Billig, *et al*,[26] for the post-surgical hip fracture patients in their case series and the 20–38 per cent prevalence reported for older persons hospitalized for a medical problem.[27,28]

• *Recovery at 12 Months:* Fifteen patients (7.8 per cent) died prior to the 12-month interview. One death occurred during the acute hospital stay and seven others before the six-month interview. This mortality experience is substantially lower than that reported previously,[5,7–9,11,29] even by studies which used restrictive selection criteria similar to ours. For example, Cobey, *et al*,[6] and Jensen, *et al*,[30] report six-month, not 12-month, mortality in comparable samples to be 7.4 per cent, and 5.8 per cent, respectively.

Descriptive statistics for the remaining recovery parameters are displayed in Table 2. Even in this selectively "healthy" sample, most individuals continued to experience a sustained decline in physical function 12 months after fracture. Fifteen individuals were permanent nursing home residents. Another 18.6 per cent were reported to be dependent on a caregiver for assistance in most activities of daily living. Indeed, while approximately 74 per cent indicated at 12 months they walked outside their home, only 21 per cent

TABLE 1—Descriptive Statistics for Selected Psychosocial Variables Measured during the Initial Post-Surgery Interview

		Excellent	Good	Fair	Poor
Global Self-Rating of Health					
% Today, day of interview		7.7	35.6	34.6	22.1
% Just before hip fracture		23.7	42.7	22.7	11.4
Personality					
Neuroticism scores:	mean 9.39 s.d. 5.3				
	standardized coefficient alpha .96				
Extraversion scores:	mean 10.04 s.d. 3.5				
	standardized coefficient alpha .92				
Depression CES-D:	less than 16		48.8%		
	16 or greater		51.2%		
	mean 16.41 s.d. 8.1				
	standardized coefficient alpha .82				
Social Connectedness:					%
Frequency of contact with "special person":			daily		54.8
			2–3 times a week		20.1
			once a week		7.3
			less than once a week		9.2
			no special person		8.7
Number of children, other relatives and friends reported being close to:			none		1.4
			1–2		10.9
			3–4		21.1
			5–6		25.7
			7 or more		40.8

TABLE 2—Descriptive Statistics for Parameters of Recovery 12 Months Following Hip Fracture Surgery among Surviving Patients

		%
Residence:	Private home: independent	62.7
	dependent	18.6
	Retirement Community	9.8
	Nursing Home	8.8
Degree of return to prefracture functional status:		
Number indicators at prefracture levels or better:	none	4.1
	1–3	27.5
	4–5	40.3
	6–7	28.1
Walking ability:	not ambulatory	9.7
	housebound	15.3
	walks in and outside home	
	uses walker	9.7
	uses quad cane	14.2
	uses straight cane	29.1
	independent, no aid	20.9
Number of self-care activities for which no assistance was received:	3	69.9
	2	14.3
	1	8.1
	none	7.7
Number of instrumental activities for which no assistance was received:	3	32.1
	2	16.3
	1	10.7
	none	40.8
Self rated health:	excellent	12.2
	good	51.5
	fair	27.6
	poor	8.2
Pain in involved hip:	none	45.9
	intermittent	44.7
	constant	9.4
Depression score (CESD):	less than eight	80.0
	eight or more	20.0
Mean	4.14 s.d. 4.3	

reported they could do so without an ambulatory aid. This compares to 81 per cent who walked outside unassisted prior to hip fracture. Moreover, only 28.1 per cent had returned to their prefracture functional status in at least six of the seven functional indicators. Twenty-three per cent had failed to regain prefracture function in at least four criteria.

In contrast to physical functioning at 12 months, the distributions on self-rated health and the CES-D indicate considerable "recovery" in these areas. The proportion rating their health as excellent or good was almost identical to those rating their health status "just before fracture" in this manner. At 12 months only 20 per cent of the sample had elevated scores, ≥ 8, on the modified CES-D scale. (Comparison of post-surgery scores on a dichotomized version of the CES-D with those on the full CES-D showed 93 per cent agreement in classification when cutpoints of ≥ 8, and ≥ 16, respectively, were used.) This proportion is similar to that reported in the total older female sample in the EPESE study[19] and substantially less than that for women in the poorest health status strata.

Predictors of Mortality

The limited number of deaths precluded an extensive investigation of the predictors of mortality. Univariate analyses involved the examination of contingency tables, and the comparison of mean values between survivors and non-survivors on continuous variables using ANOVA procedures. An increased risk of dying was observed for individuals whose cognitive status score was 3 or less, for persons rating their health, referenced to "today", as fair or poor and for those who reported having completed high school or college. There were no mean differences ($p < 0.05$) between survivors and decedents in terms of age, prefracture physical functioning, number of recorded pre-existing health problems, number of diagnoses classified as serious, number of post-surgical medical complications, fracture site, type of fracture repair, or the psychosocial variables.* Individuals who died, however, did have significantly longer lengths of stay in the acute care hospital.

TABLE 3—Adjusted Odds Ratios for Mortality in the 12 Months following Hip Fracture and 95% Confidence Intervals Computed from the Logistic Coefficients (B) and Their Standard Errors

	12-Month Mortality	
Independent Variable	AOR	95% CI
Cognitive function		
8	1	— —
7	2.09	1.35, 3.25
5	9.15	2.44, 34.36
Self-rated health		
1	1	— —
2	2.26	1.09, 4.68
3	5.09	1.19, 21.80
4	11.49	1.02, 101.79
Age in years		
60	1	— —
70	.68	.32, 1.46
80	.47	.10, 2.31
Years education		
1	1	— —
8	.93	.47, 1.83
12	.89	.40, 1.12
Number days in hospital		
14	1	— —
21	2.54	.01, 588.16
28	6.44	.53, 3493.80

Multivariate analyses used the logistic model. Variables identified as risk factors during univariate analyses were included, along with age, as predictor variables. Adjusted odds ratios (with 95 per cent confidence limits) for each variable in the model tested are shown in Table 3. Only cognitive function following surgery and self-rated health are associated with risk of death. The importance of cognitive function measured after hip fracture for risk of post-fracture mortality has been well documented, even in studies including a similar selectively healthy population.[6,8] Although not with hip fracture patients, other studies have demonstrated an association between self-rated health and mortality.[24,25]

• *Predictors of Recovery among Survivors:* A two-stage analysis limited to study survivors was conducted for five recovery indices: 1) overall 12-month physical function; 2) degree of return to prefracture functional status; 3) self-rated health; 4) pain; and 5) level of depressive symptoms. In the first stage, univariate analyses between measures obtained during the post-surgery period and the recovery indices were conducted. Selected correlation coefficients are presented in Table 4. As expected, age was strongly related to all the physical function indices with recovery more limited as age increased. Predictably, as well, initial measures of health and physical functioning were strongly associated with the com-

*Data available from author on request.

TABLE 4—Correlation Coefficients between Ordinal or Continuous Variables Measured during the Initial Post-Surgery Interview Determined from the Hospital Record and Recovery Parameters at 12 Months among the Surviving Hip Fracture Patients

	Recovery Parameters				
	12 Month Physical Function	Degree of Return to Prefracture Functional Status	Self-Rated Health	Pain	Depressive Symptoms
Age	−.38**	−.28**	.09	.08	.08
Years education	.15*	.02	−.11	−.10	−.08
Years employed	.15	.04	−.13	−.11	−.12
Number self reported medical conditions	−.31**	−.25**	.26**	.07	.21**
Walking ability	.35**	−.15*	−.08	−.02	−.14
ADL	.30**	−.05	−.05	.07	.002
IADL	.45**	.22**	−.06	−.06	−.001
Physical performance	.54**	.32**	−.29**	.02	−.31**
Physical function factor score	.56**	.25**	−.20**	.03	−.19
Mental Status	.27**	.32**	−.12	−.02	−.10
Number pre-existing medical diagnoses	−.29**	.20**	.22**	.08	.11
Number medical diagnoses considered serious	−.16*	−.08	.19**	.01	.01
Number post-surgical complications	−.18**	−.19**	.11	.04	.08
Self-rated health (before fracture)	−.30**	.28**	.47**	.09	.34**
Self-rated health (now)	−.32**	.38**	.45**	.20**	.35**
CES-D score	−.24**	.26**	.39**	.29**	.47**
Neuroticism score	−.10	−.15*	.19**	.25**	.41**
Extraversion score	.11	.06	−.19**	.08	.08
Contact "special person"	−.09	−.07	.06	.14	.02
Number close children, friends and relatives	.01	.06	−.09	−.02	−.11

*p < .05 > .01
**p < .01

parable recovery parameters. To a lesser degree these variables were related to psychosocial functioning at 12 months. With the exception of scores on the extraversion scale and the measures of social connectedness, psychosocial variables measured shortly after surgery were consistently associated with both physical functioning and psychosocial functioning at 12 months. Univariate associations between the recovery indices and treating hospital, length of hospital stay, fracture site, type of fracture repair, number of post-surgical complications, discharge location, ethnic background, prefracture living arrangement and prefracture pain were also examined. Most of these associations were not of importance.*

Multivariate analyses using OLS regression procedures were conducted. Variables shown to be related to a recovery parameter or thought to be a potential confounding variable were entered into the regression equations first; included were: age, years of education, years of employment, number of self-reported conditions, number of medical conditions reported in the hospital record, number of post surgical complications, prefracture physical function score, cognitive status score, days in the acute care hospital, and pain prior to hip fracture. Dummy variables for type of fracture repair, prefracture living arrangement, the presence of pulmonary disease and cancer, and discharge location were also entered. Each psychosocial variable was entered on a second step yielding 35 separate regression equations. Analyses were also conducted in which a best fitting model containing only statistically significant ($p < 0.05$) terms obtained using backward elimination procedures[31] was entered prior to the psychosocial variable. The results of the two analyses were identical; the former are included here. Tables 5 and 6 present the beta coefficients and standard errors for the equations evaluating the importance of post-surgical depressive symptomatology for 12-month physical function and psychosocial status respectively. Data in Table 5 confirm the previously observed importance of age, cognitive status and prefracture physical function for recovery of physical function following hip fracture. As seen in Table 6, except for pre-fracture physical function, these factors are unrelated to psychosocial status at 12 months. The data also replicate previous studies in which pre-existing medical conditions unrelated to cognitive impairments were of limited importance to recovery.[7,11] The type of hip fracture repair and post-surgical complications were also unassociated with recovery at 12 months.

In contrast, the associations between level of depressive symptoms and both the physical functioning and the psychosocial recovery indices stand out. In each case as level of depressive symptomatology following hip fracture surgery increases, outcomes 12 months later are less positive.

Comparable analyses not shown** here revealed the importance of the remaining psychosocial variables to be minimal. Specifically, the two measures of social connectedness and the measure of extraversion were unrelated to any of the recovery parameters. Self-rated health for the reference "just before you fractured your hip" and neuroticism were associated only with the psychosocial outcomes. While

**Findings from these analyses are available on request from the first author.

TABLE 5—Multiple Regression Analyses Showing the Relation between 12-month Physical Function Recovery Parameters and Selected Covariates and Depressive Symptoms (CES-D scores) Measured during the Initial Post-surgery Assessments among the Survivors at 12 Months

Independent Variables	Overall 12-Month Physical Function Beta (s.e.)	Degree Return to Prefracture Functional Status Beta (s.e.)
Age	−.034 (.008)***	−.071 (.017)***
Years educated	.030 (.017)	.063 (.037)
Years employed	.078 (.058)	.118 (.123)
Prefracture residence		
A (1= lives alone)	.036 (.138)	−.344 (.294)
B (1= lives with spouse)	.049 (.163)	−.198 (.349)
Type hip repair		
(1 = prosthesis)	.075 (.114)	.162 (.242)
Number self report conditions	−.042 (.030)	−.131 (.064)*
Number medical record		
reported conditions	−.015 (.022)	−.027 (.047)
Number post-surgical		
complication	−.063 (.087)	−.078 (.095)
Number recent hospitalizations	−.021 (.069)	−.083 (.148)
History of cancer		
(1 = yes)	−.074 (.139)	−.067 (.296)
History of pulmonary disease		
(1 = yes)	.269 (.140)	.609 (.299)*
Cognitive status	.097 (.048)*	.152 (.102)
Prefracture physical		
functioning	.465 (.062)***	.353 (.131)**
Depressive symptoms		
(CES-D score)	−.019 (.007)**	−.042 (.015)**
Constant	2.32 (.714)	10.764 (1.522)***
R	.72	.60
Adjusted R^2	.48	.31

*p.<.05 > .01
**p.<.01 > .001
***p<.001

there were associations between self-rated health for "today" and all five recovery indices, the associations between this health rating and all the recovery parameters except depressive symptomatology disappeared when the CES-D scores were added to the model.

The preceding assumes the presence of linear associations between post-surgery psychosocial variables and the recovery parameters and the absence of interaction effects. These assumptions were systematically evaluated. Analyses of residuals revealed no apparent non-linear associations. As well, none of the tested quadratic terms were statistically significant. Of the more than 100 separate analyses to evaluate the presence of interaction effects, only five achieved a p value < 0.05. Given the number of hypotheses tested, these associations may be due to chance.

Discussion

This study of the "healthier" hip fracture patient demonstrates again that older individuals who are frail have diminished reserves and are thus less able to regain losses in function following hip fracture. Impaired cognitive function as a risk factor for death or incomplete recovery of physical function is especially noteworthy. During study enrollment an experienced nurse researcher had evaluated all individuals as cognitively intact. The observed importance of cognitive status therefore reflects more subtle decrements identified by the inability to correctly respond to specific tasks on the SPMSQ such as naming the past president or subtracting from 20 by threes. Apart from the well known biomedical risk factors, this study demonstrates for the first time the importance of level of depressive symptoms as a determinant of both functional and psychosocial recovery from hip fracture.

Several methodological features of this study make it unlikely that these observed findings simply reflect study artifacts. As noted by Kasl,[13] serious methodological weaknesses of most previous studies of illness course include the inability to identify cohorts at the time of disease onset and to control for pre-existing, prognostically important aspects of health status. These problems have been minimized in this study. Hip fracture is an acute onset injury. Virtually all cases receive immediate medical attention and undergo a surgical repair. Although the initial interviews were conducted following the hip fracture and patient's responses may be affected by her convalescent status, correlations between self-rated health "just before hip fracture" and other retrospective measures such as ADL, IADL and the number of pre-existing medical conditions were substantially greater than those between these variables and self-rated health "today". Conversely, ratings of health today were more strongly correlated with other measures such as CES-D scores or cognitive status that had a post-surgery time reference. Moreover, a substantial amount of information concerning the individuals prefracture medical status was derived from her hospital record and reflected more objective assessments by the physician. Last, inclusion in the multivariate analyses of measures of pre-existing medical status, prefracture physical functioning and cognitive function minimized the possibility that the observed associations between depressive symptoms and the recovery indices simply reflect cross-sectional correlations between these prognostically important variables and depression.

The findings related to depression are in general agreement with those reported by Aneshensel, *et al.*[32] Using longitudinal data from a community dwelling adult sample, they observed illness levels to be temporally related to increased levels of psychological distress and depression to have a smaller lagged effect on subsequent illness levels. They differ, however, from those reported by Casselith, *et al*[33] who found no effect on differential survival or time to relapse among cancer patients due to any psychosocial factors studied including a measure of hopelessness. Casselith's study, however, was limited to individuals with advanced cancers where the inherent biology of the disease may override the influence of any other nonmedical factors. The results observed here were obtained in a patient cohort selected so that the individual's medical status was not likely to overshadow the effects of social or psychological factors.

While confirming the importance of level of depressive symptomatology, the study fails to demonstrate consistent associations between self-ratings of health and personality and achieved recovery. The apparent indirect pathways through depressive symptomatology for health ratings and neuroticism may reflect overlap in these psychosocial factors. Correlation coefficients between initial CES-D scores and neuroticism and health ratings "today" were .626 and .523 respectively. Although introduction of CES-D scores substantially modified the associations between these factors and the recovery parameters, the reverse did not occur. This suggests an importance of depressive symptomatology over and above that shared with either neuroticism or self-rated health. The absence of a detectable effect due to social supports may reflect inadequacies in the measurement of these variables. Only 1.4 per cent of study subjects indicated they had no person they felt very close to; thus most may

TABLE 6—Multiple Regression Analyses Showing the Relations between12-month Psychosocial Recovery Parameters and Selected Covariates and Depressive Symptoms (CES-D scores) Measured during the Initial Post-surgery Assessments among the 196 Survivors

	Self-Rated Health	Depressive Symptoms	Hip Pain
Independent Variables	Beta (s.e.)	Beta (s.e.)	Beta (s.e.)
Age	.002 (.008)	.007 (.045)	−.007 (.007)
Years educated	−.009 (.017)	−.094 (.101)	−.018 (.017)
Years employed	−.019 (.058)	.004 (.319)	−.012 (.053)
Prefracture residence			
A (1= lives alone)	−.151 (.137)	.686 (.801)	−.137 (.129)
B (1= lives with spouse)	.005 (.163)	1.721 (.909)	.228 (.151)
Type hip repair			
(1 = prosthesis)	.149 (.114)	−1.012 (.637)	.039 (.105)
Number self report conditions	.031 (.030)	−.080 (.164)	−.019 (.028)
Number post-surgical complications	−.027 (.029)	−.029 (.034)	−.008 (.006)
Number medical record reported conditions	.019 (.022)	−.060 (.123)	.017 (.021)
Number recent hospitalizations	.049 (.069)	.436 (.371)	.010 (.062)
History of cancer (1=yes)	−.074 (.139)	1.209 (.787)	.075 (.129)
History of pulmonary disease (1= yes)	−.047 (.141)	−1.939 (.760)	−.168 (.126)
Cognitive status	.072 (.048)	−.2011 (.265)	.028 (.044)
Prefracture physical functioning	−.069 (.062)	−.899 (.407)*	−.002 (.067)
Depressive symptoms (CES-D score)	.036 (.007)***	.256 (.040)***	.029 (.007)***
Constant	1.749 (.714)*	1.721 (3.932)	1.810 (.658)
R	.49	.59	.44
Adjusted R^2	.24	.28	.19

*p<.05 > .01
**p<.01 > .001
***p<.001

have exceeded a threshold distinguishing social isolates from those who are socially connected.

While the analyses reported here provide insight regarding variables that affect subsequent recovery, factors associated with the incidence of depressive symptomatology or the specific process by which recovery is affected have not been addressed. For example, we do not know if elevated post-surgery depressive symptomatology represents deterioration in affective state following hip fracture or an extension of a chronic state of depression. Likewise, we do not know if the importance of post-fracture level of depressive symptomatology is mediated through its effect on participation in a rehabilitation program, on one's ability to mobilize supportive resources or on associated biological factors that affect the capacity for recovery. Analyses of data obtained from informants and at the two- and six-month examinations are in progress and will permit further evaluation of some of these important issues.

Despite unanswered questions, the study findings emphasize the clinical relevance of routine evaluation of depression following surgery in hip fracture patients. Systematic assessment by staff nurses using one of the standard self report measures such as the CES-D or a staff rating scale such as the Raskin Scale[34] would require little time and could substantially reduce undetected depressive symptomatology. Psychiatric referral, when indicated, would permit evaluation and initiation of appropriate treatment. Although further research is required to determine the specific effect on recovery of treatment, reduction of depressive symptomatology alone is beneficial to the individual and is likely to have positive effects on those providing support and caregiving assistance as well. In light of the relatively frequent occurrence of hip fracture and the high prevalence of post-fracture elevated levels of depressive symptomatology, these benefits could be substantial.

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Delta Omega Society Announces Essay Contest: Open to Public Health Students

Delta Omega, the honorary society in public health, has announced a national essay contest on the topic of "Public Health as Social Justice." Any student currently enrolled in an accredited School of Public Health in the United States is eligible to enter. The deadline for entering contest is *April 15, 1989*.

Within the concept of "public health as social justice," the student may write about any aspect of public health in the United States or anywhere in the world. The essay may be based on empirical research, or literature review or interpretation, or on any other thoughtful approach to the subject selected. Essay length should not exceed 15 pages of text, typed double-space on 8½ × 11 inch paper. References, tables, and figures may occupy any additional five pages.

Winners will recieve cash prizes and certificates from Delta Omega attesting to their awards. First prize is $1500; second prize, $1000; third prize, $500. All three winners will be invited to present their papers during the 117th annual meeting of the American Public Health Association in Chicago, Illinois, October 22–26, 1989. Delta Omega will also contribute to the travel expenses of essay winners.

Selection of the winning essays will be made by a Delta Omega Committee of faculty members from at least three accredited Schools of Public Health. Send all essays by first class mail to:

Professor Henrik Blum
Chair, Delta Omega Essay Contest
School of Public Health
University of California
Berkeley, CA 94720

Preventing Alcohol-Impaired Driving through Community Self-Regulation Training

JOHN K. WORDEN, PHD, BRIAN S. FLYNN, SCD, DARWIN G. MERRILL, MS, JULIAN A. WALLER, MD, MPH, AND LARRY D. HAUGH, PHD

Abstract: A community education program was designed to train the individual drinker to self-regulate his or her blood-alcohol concentration (BAC) below a level of impairment (.05 g/dl or 11 mmol/L). Drink calculators (cardboard wheels and wallet cards) were disseminated to customers of bars and licensed beverage outlets; bartenders and counter clerks were trained to demonstrate use of the calculators and demonstrations were presented in television spots. Program components were evaluated in three matched Vermont communities, one receiving the full community education program, one receiving the TV spots only, and one serving as control. After six months of intervention, a roadside survey of nighttime drivers (N=892) indicated 5.3 per cent fewer drivers with BACs above 0.05 g/dl in the community program group and 1.0 per cent fewer in the TV-only group compared to the control group; however, substantially fewer drivers were found above .00 BAC in either program community than in the control. Drivers reporting heavy drinking and youthful drivers both indicated higher utilization of the materials than did other drivers. Although limited in scale and duration, this study suggests that a community education program can be effective in preventing alcohol-impaired driving. (*Am J Public Health* 1989; 79:287–290).

Introduction

Alcohol-impaired driving remains the leading human cause of highway deaths, despite repeated efforts to deter this behavior through a variety of interventions. In the past, public education has been employed largely to raise awareness of alcohol and highway safety issues and to promote moral persuasion, while most direct interventions to reduce alcohol-impaired driving have focused on external regulation through legislation and enforcement.[1–5] Server interventions show potential for prevention although they have not been evaluated with driving populations.[6,7]

In this study, we have evaluated a community education program which was designed to provide the individual drinker with a cognitive method of self-regulating his or her blood-alcohol level through the use of "drink calculators" which were distributed and demonstrated at licensed alcoholic beverage outlets. In Vermont, such calculators have been used sporadically for a number of years, particularly in school driver education courses, but their use in a concentrated community-wide education program has not been evaluated previously.

Methods

Educational Program

The community education program included: a series of meetings with community leaders; a system through which drink calculators were distributed and demonstrated to community members through licensed alcoholic beverage outlets; and broadcast of public service television spots in which use of drink calculators was demonstrated. The full program was implemented in a single Vermont community with a target population of about 4,000 persons of legal age for both drinking and driving.

Drink calculators are inexpensive devices which provide estimates of blood-alcohol concentrations (BAC) for individuals by body weight and rate of alcohol consumption. Persons using the calculators can estimate the number of drinks they could consume in a given period of time without exceeding the Vermont legal driving limit of 0.10 g/dl BAC (22 mmol/L); estimates are also provided for the number of drinks likely to produce a BAC of 0.05g/dl (11mmol/L), at which impairment of driving skills begins to be perceptible for most people. A drink is defined as one ounce of 100 proof liquor, 12 ounces of beer, or 4 ounces of table wine. Two types of drink calculators were used. Drink calculator wheels are 3″×4″ cards with a rotating wheel mounted. Using the wheel to match body weight and number of drinks consumed provides an estimate of BAC level in a window; a second window provides a time adjustment if these drinks were consumed in a period of more than one hour. The second type of calculator was a simplified wallet-sized card version of this device. Directions for use and motivational messages were printed on the drink calculators.

The distribution-demonstration system for drink calculators was implemented by visits to each retail outlet every four weeks on average throughout the six-month intervention period; visits were scheduled to cover different shifts and were more frequent at the beginning of the period, less frequent later. Service personnel were asked to rehearse demonstration skills and any difficulties were addressed. They were urged to demonstrate and distribute the calculators to customers. Brightly-colored countertop displays which held the calculator wheels were installed prominently at each location.

Logs of these contacts recorded 150 visits to retail outlets. Fifty-three customer service personnel were trained directly in use of the calculator, who, in turn, trained a number of other customer service personnel. Comments from the outlet workers and owners reflected generally good acceptance throughout the implementation period. The calculators were distributed through all 25 retail alcoholic beverage outlets identified in the program community, including: convenience stores (6), restaurants (6), private clubs (5), filling stations (3), lounges (2), supermarkets (2), and one

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state liquor store. No licensees refused to participate. About 3,000 drink calculator wheels were distributed through countertop displays, and about 3,000 wallet cards were distributed with purchases of alcoholic beverage. Excess distribution above the target population of 4,000 persons can be attributed to visitor traffic and multiple liquor sales per person.

In addition to face-to-face instruction, use of drink calculators was demonstrated in a series of television public service announcements. A 60-second message featured a well-known race car driver, who was also a local resident, demonstrating use of a drink calculator to his racing crew. Two 30-second messages featured demonstration by popular local bartenders. These messages were broadcast an average of 20 times during each week of the six-month intervention period, primarily during late afternoon and late night hours. The television station was a local network affiliate, with about a 20 per cent share of the audience, whose management undertook the project as a major public service initiative.

Evaluation

The impact of the educational program was evaluated through a controlled quasi-experiment in three Vermont communities. The first community (A) received the full educational program; the second community (B) received broadcasts of the drink calculator television spots but no other element of the educational program; the third community (C) was a measurement-only control.

The three communities were selected from two Vermont counties which had, for the most recent five years, the highest average annual automotive fatality rates (0.33 deaths per 1,000 residents). One of these counties was served by a television station which could not be seen in the second county. Within the former county, the two largest communities were found to provide good demographic matches (see Table 1); the smaller community was designated to receive the full educational program (A) and the larger received television spots only (B). One of two equivalently sized communities in the second county providing the better demographic match was chosen as the measurement-only control site. The two program communities were located about 20 miles apart. The measurement-only community was in a non-adjacent county 80-90 miles from the program communities; a mountain range separates the two areas.

Roadside surveys were conducted in the three communities during April and May 1983 on Friday and Saturday nights between 9 pm and 3 am; police officers stopped all drivers at previously designated locations and visually checked them for signs of drinking. Vermont drivers (N=1,332) were asked to participate in a brief survey administered by the officers; there were no refusals. The officers were not aware of the study design. The survey asked about usual frequency and quantity of alcohol consumption and knowledge of the number of drinks in an hour to reach a BAC of .10 g/dl (22 mmol/L).

TABLE 2—Reported Usual Quantities of Alcohol Consumed by Drivers in Each Study Group at Follow-up Survey

Study Group	Usual Number of Drinks*		
	None	One or Two	Three or More
	%	%	%
Control (N = 311)	19.5	58.8	21.7
TV Only (N = 377)	19.2	60.1	20.7
Community Program (N = 132)	20.5	57.0	22.5

*Number of drinks usually consumed in a one-hour period.

For consistency with program exposure and to obtain the maximum number of respondents, survey participants were assigned to study group by town of residence regardless of the interview location. Participants living outside the boundaries of the community receiving the full educational program but within range of the television station which broadcast the spots were included in the TV-only group; participants residing in other Vermont towns outside of the range of the television signal were included in the measurement-only group.

In October and November 1983, after completion of the six-month educational program, a more comprehensive series of roadside surveys was conducted at equivalent times on weekend nights. Police officers stopped all traffic and, except for a few drivers considered by police to be legally impaired, all Vermont drivers were passed on to a separate team of interviewers. These drivers were assured that, if found to be impaired during BAC measurements given voluntarily, no penalties would be exacted. If drivers and all other occupants of their cars were found to be impaired or otherwise unable to drive they were driven home by members of the survey team. The interview teams collected survey data, including BACs, from all consenting Vermont drivers; BACs of drivers processed by police were also included in the analysis. Of 1,009 eligible drivers, 927 (91.9 per cent) consented to be interviewed; 881 of these 927 (95% per cent) consented to have their BACs measured. Refusal rates by study area were nearly identical. Sixty-nine per cent of the drivers were males, and 53 per cent were age 30 and under; these distributions were nearly identical among the three study areas.

Results

In the baseline survey, there were no differences in the usual drinking patterns reported by drivers or in their estimates of the number of drinks which they could consume within a one-hour period and not be considered legally impaired for driving. Police officers reported that 16.8 per cent of the surveyed Vermont drivers had been drinking: 16.3

TABLE 1—Population Characteristics of Study Communities

Study Area	Education High School and Above	Income Median Household	Age: (Proportion 16–29) Male	Age: (Proportion 16–29) Female	Population (1980 Census)
	%		%	%	
Community A	19.5	$13,086	11.3	12.3	5538
Community B	19.7	$13,326	10.6	12.6	11886
Community C	19.8	$12,723	12.8	13.4	5068

TABLE 3—Blood-Alcohol Concentrations (BAC) for Drivers in Study Groups (Follow-up Survey)

BAC	Control (N = 333)	TV Only (N = 405)	Community Program (N = 154)
	%	%	%
0	74.5	80.7	84.4
.01–.04	14.4	9.1	9.7
.05–.09	8.1	5.4	5.2
.10 and Above	3.0	4.7	0.6

per cent were reported in the community program group, 19.3 per cent in the TV-only group, and 14.2 per cent in the control group. As indicated in Table 2, reported drinking patterns also were found to be similar among all three study groups in the follow-up survey.

Program effects were assessed in the follow-up survey using the breath sample measurements of BAC as indicated in Table 3. Only 5.8 per cent of the drivers in the community program had BACs above the level of impairment (.05 g/dl or 11 mmol/L), compared to 10.1 per cent in the TV-only group and 11.1 per cent in the controls. The odds ratio for the community program group staying below the level of impairment relative to the control group was 0.51 (95% CI: 0.24, 1.10); for the TV-only group the odds ratio for BACs below the level of impairment was 0.91 (95% CI: 0.56, 1.45) compared to controls. The odds ratio for a BAC of .00 for the community program group relative to the control group was 0.54 (95% CI: 0.33, 0.91); for the TV-only group this odds ratio was 0.74 (95% CI: 0.49, 0.99). Results from Jonckheere's distribution-free test for ordered alternatives[8] based on recorded BAC levels supported the hypothesized ordering of effects ($p=.03$).

Mediating variables were found to be consistent with these results. As indicated in Table 4, 47 per cent of the drivers residing in the town receiving the full community program said that they had used a drink calculator wheel, 59 per cent said they still possessed a drink calculator wheel at the time of the follow-up survey, 56 per cent recalled seeing the countertop display, and 81 per cent reported seeing the drink calculator wheel demonstrated on television, compared with substantially lower levels of activity in the other two groups.

Driver recall data also indicated that persons at highest risk of alcohol-impaired driving, because they reported usually drinking three or more drinks at a time (i.e., in "an hour or so"), were most likely to be exposed to the drink calculator wheel cards. Table 5 shows higher levels of exposure to program components for high-risk drivers in the town receiving the full community program compared to drivers at lower risk residing in the same town. Similar patterns also exist for high- and low-risk drivers in the TV-only and control groups.

A higher level of exposure to the full range of educational approaches was also reported by youthful drivers (30 years of age and younger) compared to older drivers residing in the town receiving the full community program, as indicated in Table 6. Among youthful and older drivers in the TV-only and control groups, similar exposure patterns were found for drink calculator wheel use. Youthful drivers were more likely to recall the TV spots in the TV-only area than older drivers; no differences were found with regard to the other educational approaches.

TABLE 5—Reported Use of Drink Calculators by Heavy, Light, and Non-drinkers in Community Program Group

Drinker Type	Used Wheel	Used Card
	%	%
Heavy Drinkers* (N = 34)	62	45
Light Drinkers (N = 87)	52	27
Non-Drinkers (N = 31)	23	10

*Reported usually drinking three or more drinks at a time (i.e., in "an hour or so").

Discussion

Questions have been raised about the appropriateness of using drink calculators. For example:

- They are calibrated primarily for a young male population and may give BAC estimates per drink as much as 15 per cent too low for females and the elderly who have different body fat distribution.[9]
- While the dose-BAC relationship may be accurate on average for young males, there is enough personal variation in alcohol metabolism so that some individuals who carefully follow the calculator may still exceed the legal limit if their goal is simply to stay legal.[9]
- The calculators are based on an assumption of a standard "drink" containing about 18 cc (0.6 oz) of pure alcohol. But, as served, many drinks of liquor contain much larger quantities of alcohol, and the distinction between a "standard" drink and a "real" drink for those consuming hard liquor may be too subtle for much of the intended audience.
- Recent reanalyses of epidemiologic and laboratory data indicate that for many people alcohol-related impairment actually begins below 0.5 g/dl (11 mmol/L), although it does not become appreciable until levels of .08 g/dl (18 mmol/L) or higher are reached.[10]

These arguments, although cogent, are insufficient to warrant abandonment of the drink calculator as a useful tool for self-regulation of drinking. Although more women are drinking and drinking heavily as evidenced by crash fatality trends, young males still comprise the vast majority of those who drink heavily and get into trouble with alcohol-impaired driving. These individuals commonly drink enough to reach the maximum legal BAC, and often go far beyond it. If the goal were simply to get these people down to the legal limit, many of them would have to cut their alcohol consumption in half. But the goal of the drink calculator educational program

TABLE 4—Reported Exposure to Drink Calculator Wheels and Cards by Drivers in Each Study Group

Study Group	Used Wheel	Have Wheel	Saw Wheel on Counter	Saw Wheel on TV	Used Card
	%	%	%	%	%
Control (N = 337)	31	29	10	22	10
TV Only (N = 405)	24	36	17	50	9
Community Program (N = 154)	47	59	56	81	28

TABLE 6—Reported Exposure to Drink Calculator Wheels by Young and Older Drivers in the Community Program Group

Age (years)	Used Wheel	Have Wheel	Saw Wheel on Counter	Saw Wheel on TV
	%	%	%	%
30 and Under (N = 93)	61	68	68	88
Over 30 (N = 66)	27	45	38	71

was to convince individuals that a more sensible approach is to stay below the level of .05g/dl (11 mmol/L).

Roadside survey methods for research purposes, such as those used in this study, have been examined by several courts in relation to Fourth Amendment issues of search and seizure. One opinion by a judge of the US District Court of Minnesota (*Stark vs Perpich, et al,* 1984) suggested methods other than those described in this study. Because US courts are still considering these issues, we as authors feel that we are not in a position to say which of these methods is most appropriate for use in future roadside research.

Limitations of the present study include a six-month intervention and a posttest-only measurement of BAC using three matched communities. Fortunately, there were no differences in the age, gender, and reported drinking patterns of respondents among the three study groups. Proportions of drivers observed by police officers to have been drinking were also similar among study groups, although visual observation is generally imprecise compared to chemical measures.

To maximize usable data collected from nighttime drivers in a small rural state, all Vermont drivers stopped in roadside surveys in the study communities were included by assigning them to study groups based on town of residence and the level of program exposure reaching each town. The modest size of the community receiving the full educational program restricted the number of drivers available for roadside surveys in that area and resulted in a smaller sample than was expected for statistical analysis; results must be viewed with this limitation in mind.

Media exposure was distinctly partitioned by the Green Mountain range to exclude a vast majority of Vermont towns; reported exposure to the TV spots in the control group (22 per cent) was similar to that noted in previous research,[11] and can be attributed to confusion with similar media materials, agreement bias and, to a small extent, true exposure although mobility between control and program communities is limited. It is also interesting to note the increased awareness of the TV campaign in the community program group (81 per cent) compared to the TV-only group (50 per cent) while the channel availability is similar between areas; this may be an example of the supportive roles played by interventions on both the interpersonal and mass media communication levels observed in other intervention studies.[1,2]

The present study is based on the distribution of drink calculators and the training of customer service personnel at alcohol beverage outlets throughout the community. This study has resulted in the development of a community education program involving relatively low cost in personnel and material. The total estimated cost of the six-month program was $3,325: TV-spot production, $1800; materials, $725; personnel and travel, $800.

Results suggest that the community program with mass media support was more effective in preventing alcohol-impaired driving than were the media messages alone, but that the media messages alone also produced a general behavioral effect of deterring the consumption of any alcohol before driving. The primary distinction in results between the groups receiving the full community program and the media intervention alone appeared to be in the proportion of drivers having BACs above the state's legal limit: 0.6 per cent in the community program compared to 4.7 per cent in the TV-only group. This result tends to underscore the role of skills training, since both approaches emphasized the importance of staying below the state's legal limit (as well as the .05 g/dl or 11 mmol/L level of impairment), but the hands-on experience of operating the drink calculators with personal corrective feedback may have served to reinforce the message. Anecdotal reports also reflected social support in the community through the sharing of drink calculators and discussion of how to use them. This self-regulation approach could be included as a component of server interventions described by Mosher.[6]

Most importantly, this type of intervention has stimulated self-regulation by the individual drinker without direct coercion by outside authority. In addition, the known presence of enforcement was an important positive incentive to persons living in all of the study areas. However, this education program has had the distinct advantages of being truly preventive of hazardous behavior and not requiring corrective action by police or the courts. It also was designed to give the individual the skills and social support to develop long-term positive habits which can eliminate the personal risk of driving while impaired by alcohol.

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Traffic Safety Effects of Sobriety Checkpoints and Other Local DWI Programs in New Jersey

David Levy, PhD, Dennis Shea, MA, and Peter Asch, PhD

Abstract: We examined the efficacy of three Driving While Intoxicated (DWI) programs in New Jersey from 1980 through 1985, using covariance analysis of county data. Road blocks, the major component of the Strike Force program, were associated with a drop of 10–15 per cent in the single vehicle nighttime crash rate and showed a relatively stable effect over time. DWI Task Force, an education program, was associated with a 6–10 per cent total decline in the crash rate and declining impact over time. SOBER, another education program, was associated with a small effect in the first year and little or no effect thereafter. (*Am J Public Health* 1989; 79:291–293.)

Introduction

Most states and communities in recent years have adopted a variety of programs in an attempt to reduce alcohol-related accidents and fatalities on the highways and roads. Programs have been designed to educate drivers, elicit community support, and strengthen preventive and punitive programs. One of the more controversial efforts is the roadblock or sobriety checkpoint programs, in which drivers are stopped to determine whether they have been drinking while driving. These programs are currently in use in a number of states and under active debate in a number of others.

Since the United States Supreme Court's decision in *Delaware v. Prouse*,[1] checkpoint programs have been subject to the requirement that effectiveness be demonstrated. The Court noted that the "the permissibility of a particular enforcement practice is judged by balancing its intrusion on the individual's Fourth Amendment interest against its promotion of legitimate governmental interest."[2] Yet, little attention has been directed toward evaluating the effectiveness of such policies. Ross, *et al*,[3] found very limited effects in Scandinavian nations, concluding that "there is no evidence of important deterrent consequences." Mercer found that road checks alone were not associated with injury reductions in British Columbia.[4] Two unpublished studies have reported statistically significant reductions of checkpoint programs on traffic deaths and injuries.[5,6]

We report here the results of some studies of the effectiveness of DWI (Driving While Intoxicated) programs in New Jersey. Effectiveness is defined in terms of traffic crash experience, and special attention is devoted to the effects of the Strike Force (checkpoint) program.

Local DWI Programs in New Jersey

Since the early 1980s, state and local programs have been vigorously pursued to reduce the drinking and driving problem in New Jersey. Local level programs include two educational and community support efforts—SOBER (Stay Off the Bottle, Enjoy the Road) and DWITF (Driving-While-Intoxicated Task Force)—and an enforcement and deterrence program, Strike Force (roadblocks). These programs are administered at the county level, with supervision and funding from the state and federal government. Starting dates for programs by county are shown in the Appendix.

The Strike Force program provides overtime funding to police departments on a county-wide basis and uses a system of random roadside checkpoints to examine drivers for possible intoxication. Police direct traffic onto a single lane, where officers speak to the driver and provide drunk driving information materials. Drivers who appear to have been drinking are directed to an area off the roadway for further screening (psychomotor and breath tests).

The checkpoint sites and times are determined by police personnel based on prior analysis of accident and arrest data. Operations are usually conducted on weekend nights and are moved to different sites. Table 1 provides the number of hours worked by police, total costs, number of vehicles stopped, and number of arrests for Bergen County, the first county in New Jersey to implement the Strike Force program and the pilot for other counties. County-wide programs have been supplemented by state roadblocks on state roadways. These were conducted throughout New Jersey from early 1983 in a manner similar to local roadblocks.

Although the Strike Force programs remain small in terms of resources, their impact on public consciousness is important. The county checkpoints have generally been implemented after a SOBER and Task Force program, and the checkpoint programs themselves were publicized on radio and television. The sites are unannounced, but visibility is a prime consideration. Checkpoint trailers with banners were conspicuously parked in the participating county and driven in major areas when not in use. Surveys conducted for the state found awareness of them is close to universal among drivers.

An important part of the deterrence strategy is the informational and consciousness-raising programs, which not only educate citizens but also inform them of enforcement efforts. At a county level, the SOBER and DWITF programs employ pamphlets, public service announcements, seminars for students, and other publicity campaigns about the dangers of drinking and driving, and are directed at educating citizens and eliciting community support. These programs are supplemented by state level efforts and publicity through both the electronic and print media. In 1984, the

TABLE 1—Summary of Strike Force Activities in Bergen County, NJ, 1983–85

Year	Months	Hours	Cost Worked	Vehicles Stopped	Arrests
1983	8	2340	$ 46,800	29388	393
1984	12	5332	$129,660	45612	456
1985	12	4805	$110,510	38833	258

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SOBER programs in New Jersey (implemented in all but one county by October) had authorized funding of $173,000 and involved 460 press releases in county newspapers, 100 public service announcements on local radio stations, $365,000 worth of advertising on all media, distribution of one million pieces of literature, and 50 billboard advertisements. In that year, DWI Task Force had 640 members in 11 counties, provided two conferences as well as DWI training for police speakers, and was involved in data gathering, lobbying, and monitoring of courts. In 1985, most of the activities of both programs increased by about 10 to 20 per cent.

Methods

Total fatality and crash rates provide a broad measure of traffic safety. To gauge more accurately the effects on alcohol-related events, we focus specifically on single-vehicle nighttime (between 6 pm and 5:59 am) crashes and non-pedestrian fatalities involving a motor vehicle. These measures are regarded as a good proxy for alcohol-related events because of their objectivity, i.e., there is virtually no judgment involved in classifying a single-vehicle nighttime event.[7,8] The drawback of the measure lies in its broadness. Although a preponderance (approximately 70 per cent) of single-vehicle nighttime crashes and fatalities are believed to be alcohol-related, the category includes some events that have nothing to do with alcohol consumption and excludes over events that are alcohol-related. The data were obtained from the New Jersey Office of Highway Safety and are also available from the National Highway Traffic Safety Administration. The fatality data were deflated by total population.

We examined general trends in New Jersey relative to the United States. Using covariance analysis, we also estimated the effects of the different traffic safety programs and other factors on the number of crashes per capita in a particular county. Pooling cross-sectional county data over time, we examined the behavior of traffic safety variables for each of New Jersey's 21 counties over the six-year period, 1980–85, for a sample of 126.

The dependent variable in the covariance analysis was the logarithm of the number of nighttime single-vehicle crashes deflated by county population in the observation year. Upon estimating equations with total fatality and blood alcohol concentration (BAC)-related fatality rates as the dependent variables, all program coefficients had the expected negative signs but were often statistically insignificant. The reason may be that the numbers of fatalities, measured on an annual county basis, are quite small and show considerable volatility over time. Also, BAC tests are not conducted on all fatal victims.

The independent variables in the covariance equations consisted of dummy (i.e., indicator) variables for each county, year, and intervention. The county variables control for the various demographic, highway, and geographic elements specific to a county. The time dummy variables for each year account for general trends in crash patterns. A separate variable was included for each of the three interventions—SOBER, DWITF, and Strike Force—to indicate the presence or absence of each program in each county and year. The value of each variable is zero unless the program was in effect in a given county and year. Then, the value of the variable was defined by the weighting system described below times the percentage of the year in which the program was in effect.

The pattern of influence over time of the programs on the crash rate may vary, e.g., with a significant impact in their first year of operation and then declining effects over time or, alternatively, a slower "start-up" period followed by a greater impact when drivers realize the new efforts at DWI enforcement. We first added dummy variables to the equation for first-year effects, but this yielded a high degree of multicollinearity. Consequently, the first year dummy variables were dropped, but the results from the analysis, along with an examination of the results from a time series intervention analysis that we do not report here,* were used to distinguish a series of weights for the number of years in which the program was in existence.[9,10] The respective weights for the first, second, and third years were 1.0, 0.05, and 0.25 for SOBER, 0.5, 1.0, and 0.5 for DWITF, and 0.5, 1.0, and 0.75 for Strike Force. The weights on the SOBER variable reflect an abrupt first-year effect that declines sharply thereafter. In contrast, the Strike Force Program effect was gradual, increasing and remaining relatively stable after the first year, whereas the DWI Task Force effect occurred gradually but declined more sharply after the second year.

We included lagged values of the dependent crash rate variable in some equations to control for the "inherent" safety characteristics of various counties and for correlation in the error term over time for each county. The problem of uneven error terms arising from populations varying across counties and over time (i.e., heteroscedasticity) is treated by a method adapted from Cook and Tauchen.[11]

Results

Preliminary evidence on traffic fatality and crash rates is provided in Table 2. In New Jersey between 1980 and 1985, total crash rates rose by 8 per cent while total fatality rates fell by 16.2 per cent. During the same time period in New Jersey, single-vehicle crash rates fell by 18.6 per cent and fatality rates fell by 27.9 per cent, both falling considerably more than total fatality rates. For the United States, consistent data were available for fatality rates, but not crash rates. For the US between 1980 and 1985, all fatality rates fell by 18.3 per cent, 2.1 per cent more than the New Jersey rates, and single-vehicle fatality rates fell by 24.5 per cent, 3.4 per

*A Box-Jenkins analysis, allowing for a flexible shape of the intervention variable, was conducted for Bergen County, using monthly data for 1980–85. The SOBER and DWITF programs exhibited declining effects over time, while the effects of the Strike Force program were found to increase over time.

TABLE 2—Fatality and Crash Rates, by Year, per 1000 Population

Year	NJ Total Crash Rate	NJ Total Fatality Rate	NJ Single Vehicle Nighttime Crash Rate	NJ Single Vehicle Nighttime Fatality Rate	US Total Fatality Rate	US Single Vehicle Nighttime Fatality Rate
1980	29.9	0.1521	4.13	0.0649	0.2248	0.0871
1981	29.6	0.1569	4.03	0.0583	0.2147	0.0813
1982	29.9	0.1428	3.91	0.0562	0.1894	0.0748
1983	28.6	0.1248	3.35	0.0493	0.1818	0.0694
1984	29.7	0.1227	3.25	0.0404	0.1872	0.0696
1985	32.3	0.1275	3.36	0.0468	0.1836	0.0657
% Change 1980–85	8.03	−16.2	−18.6	−27.9	−18.3	−24.5

cent less than the comparable New Jersey rates. The fall in New Jersey's single-vehicle nighttime rates relative to the state's all-fatality rate and relative to the US single-vehicle nighttime fatality rate would suggest that New Jersey was effective at deterring alcohol-involved traffic crashes. However, a number of other influences may have contributed, such as other statewide policy changes in New Jersey (including two increases in the minimum legal drinking age and stricter court penalties for drinking and driving).

Equations explaining single-vehicle nighttime crashes at the county level are presented in Table 3. The county and time dummy variables are not reported here (since they are not of direct interest) but these variables were in almost all cases significant at standard levels. Equation #1 (without the lagged dependent variable) and equation #2 (with the lagged dependent variable) have high explanatory power (in terms of the R-squared), and all program variables have the expected negative sign. The coefficient for the Strike Force variable indicates that the program at its peak reduced the single-vehicle nighttime crash rate by between 10 per cent [in equation #1, with a 95% confidence interval (CI) = −0.175, −0.015] and 15 per cent [in equation #2, with 95% CI = −0.25, −0.040]. Based on the weighting scheme, the peak effect is reached in the second year and then gradually declines. The DWITF program reduces the crash rate between 6 per cent [95% CI = −0.162, −0.33] and 10 per cent [95% CI = −0.175, 0.053], but this effect declines more quickly over time. The effect of SOBER is near zero and insignificant, but approaches significance using a one-tailed test in equation #1.

Discussion

The estimated program effects, particularly for the checkpoint program, may be understated since programs in one county may have "spill-over effects" in terms of driving behavior in other counties. Not only do the estimated program effects fail to directly take into account any spillover in terms of crashes avoided in nearby counties, but may also bias downward the estimated effect in the county implementing the program. Drivers may avoid drinking and driving in a county without a checkpoint program, because they must pass through a participating county to get to their destination. Furthermore, state checkpoints at the time of the county programs may have lowered crash rates statewide.

TABLE 3—Covariance Analysis of Single-Vehicle Nighttime Crash Rates[a,b,c]

Independent Variables	Equation 1	Equation 2
Lagged Dependent Variable		3.37 (1.48)
SOBER	−0.025 (−1.223)	−0.009 (−0.246)
DWITF	−0.096* (−2.85)	−0.061 (−1.06)
Strike Force	−0.095* (−2.33)	−0.143* (−2.71)
Adjusted R-Square	0.98	0.93
Degrees of Freedom	97	96

a) Log form, corrected for heteroscedasticity (Cook-Tauchen[11] correction).
b) By county, 1980–85.
c) t-statistics in parentheses.
*indicates statistically significant from zero at the 0.01 level (one-tailed test).

The findings suggest that a checkpoint or "roadblock" program such as New Jersey's Strike Force reduces single-vehicle nighttime crashes. Besides the results reported here, separate time-series analyses for one of the demonstration counties in New Jersey strongly support this conclusion, indicating program effects of greater than a 20 per cent reduction in the single-vehicle nighttime accident rate at their peak. The effects of programs oriented toward publicity and education about the dangers of drinking and driving are less impressive; even here, however, we suspect that judicious, perhaps sporadic, use of such campaigns may have temporary beneficial impacts. However, it may well be that the effectiveness of the checkpoint program itself depends partly on the publicity that accompanies its implementation.[4]

APPENDIX
Local DWI Programs in New Jersey by County and Start Dates

County	SOBER	DWITF	Strike Force
Atlantic	10/82	10/85	
Bergen	10/82	1/83	5/83
Burlington	10/84		
Camden	10/82		
Cape May	10/84	1/85	
Cumberland	10/82	9/16/84	
Essex	10/82		
Gloucester	10/82		
Hudson	5/86	5/85	
Hunterdon	10/81	3/86	
Mercer	10/82	3/84	8/84
Middlesex	10/83	6/84	11/84
Monmouth	10/80	1/83	5/84
Morris	10/80	5/82	5/84
Ocean	10/81	1/86	
Passaic	10/84	1/83	3/84
Salem	10/83		
Somerset	10/84	5/85	
Sussex	10/82	8/83	
Union	10/82	1/85	
Warren	10/84	1/85	

ACKNOWLEDGMENTS

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Blood Alcohol Tests, Prevalence of Involvement, and Outcomes Following Brain Injury

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Abstract: We collected data on all residents of San Diego County, California who were hospitalized for or died from a brain injury in 1981. The objectives were to assess the frequency of blood alcohol concentration (BAC) testing and the associations of BAC prevalence with the external cause of the brain injury and case outcome. We found that high BAC levels were most frequent among brain-injured subjects between the ages of 25 and 44 and among those subjects involved in motor vehicle crashes and assaults. Contrary to expectations, injury severity and hospital mortality were inversely related to BAC level, controlling for other predictors. We believe that these inverse associations might be due to differential rates of BAC testing by severity. Among brain-injured survivors with more severe injuries, however, we found that BAC level was positively associated with the prevalence of physician-diagnosed neurological impairment at discharge and with the length of hospitalization. (*Am J Public Health* 1989; 79:294–299.)

Introduction

The scientific literature on alcohol and motor vehicle crashes points to high risk of crash involvement when alcohol is consumed.[1] However, less information is available on many questions about alcohol use and injuries. For example, which subgroups of injured motorists are blood-alcohol-tested and, by extension, are the published findings representative of all those injured?

A related question is the relation of alcohol to the occurrence of some types of nonmotor-vehicle-related injuries. Although the association between blood alcohol concentration (BAC) and risk of injury is well established for motor vehicle crashes,[1] general aviation crashes, drownings, snowmobile crashes, and violence,[2–4] there are less well established linkages between alcohol and injuries from falls and from injuries at work or during some recreational activities.

A third question involves the role of alcohol in the outcome of injury. Animal studies[5–15] have demonstrated a variety of physiological effects of alcohol (ethanol) on injuries. In humans, however, the findings are equivocal. Some reports[16,17] suggest that intoxicated persons are somehow protected from serious injury; other studies suggest that intoxication leads to more severe injuries.[18–20] An important public health aspect of these findings is whether the physician's ability to establish accurate diagnoses for persons with neurological injuries[21–25] is compromised by the presence of alcohol.

This report describes BAC testing among a cohort of San Diego County, California residents who were brain-injured in 1981 and examines the rate of BAC testing, the proportion of positive BAC test results, the relation of BAC level to brain injuries from various external causes, and the relation between BAC level and injury outcome.

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Methods

Members of the cohort were residents of San Diego County, California who were hospitalized (or died) in 1981 due to brain injury. Specific details of the study design, study region, case ascertainment, and methodology are reported in detail elsewhere,[26–28] and certain relevant points are summarized here.

For purposes of this study, brain injury is physical damage to, or functional impairment of, the cranial contents from acute mechanical energy exchange exclusive of birth trauma. Information was ascertained on persons with autopsy evidence of brain injury who died at the injury scene or were dead on arrival at an emergency room and on persons admitted to a hospital with physician-diagnosed brain injury.

Brain-injury cases were identified from emergency room and inpatient hospital records of acute care general hospitals in the county; all coroners' records in San Diego County and the adjoining counties of Imperial, Orange, and Riverside; all death certificates (regardless of cause or place of death) for residents of San Diego County; a survey of all nursing homes and extended care facilities in San Diego County; and reviews of the nine major hospitals in the three counties bordering San Diego County. Autopsy and coroners' records were obtained for all relevant fatalities.

For people surviving until hospital admission, the Glasgow Coma Scale (GCS) was used to assess the level of brain-injury severity.[29] This scale was used consistently in all San Diego County emergency rooms during 1981. For a few patients, a verbal response was not recorded (e.g., because of intubation), and it was assumed to be unimpaired if eye and motor response were unimpaired. For a very small number (less than 1 per cent) of severely injured patients for whom the GCS was not assessable, the medical chart was examined thoroughly and all relevant clinical information was used to judge brain-injury severity. The GCS used for these analyses was measured upon arrival at the emergency room.

Patients admitted with a GCS of 8 or lower were defined as *severe* cases; those with a GCS of 9 or greater were defined as *moderate* if they had a hospital stay of at least 48 hours and either a GCS below 13, an abnormal computerized axial tomography (CAT) scan, or brain surgery. All other cases were termed *mild*. These categories and criteria are consistent with those suggested by Jennett and Teasdale[29] and Levin, *et al.*[30]

Overall injury severity was defined using the 1980 revision of the Abbreviated Injury Scale (AIS) which assigns

a severity level from 1 to 6 to each injury.[31] The Injury Severity Score (ISS) assigns an overall measure of injury severity based on AIS scores.[32,33] Injuries are assigned to six body regions (face, head and neck, chest, abdomen, extremities, and general), and the ISS is calculated as the sum of the squares of the highest AIS values for three separate regions. A patient's ISS is a good measure of the risk of death from injuries, and this predictive power is improved by age-adjustment.[32,33]

Rates of neurological limitations were derived for all moderately or severely brain-injured persons admitted to a hospital. These rates were adjusted for age, gender, and injury severity at the time of emergency room treatment. Neurological limitations included only physician-diagnosed deficits or limitations at the time of hospital discharge.

Persons under the age of 15 have been excluded from the analysis because they were not generally tested for blood alcohol by either hospitals or coroners in 1981.

Blood alcohol samples were generally obtained while the person was in the emergency room, and results were abstracted from the hospital record or the coroner's report and were not based on police reports. Most hospitals and coroners used a gas chromatographic method to determine blood alcohol concentration.

The effects of BAC and other factors on injury severity and hospital mortality were estimated and tested with multiple logistic regression.[34]

Results

Of the 3,358 new cases of brain injury identified, 709 were under the age of 15 and excluded from the analyses.

BAC Testing

Forty-four per cent of 2,649 brain-injured persons age 15 and older were BAC tested. The overall BAC testing rate was higher for males (49 per cent) than for females (30 per cent) and varied by age. BAC testing was lowest in young adults (32 per cent) and those over age 64 (36 per cent), and it was highest for those aged 40–49 and 60–64 (52 per cent and 55 per cent, respectively).

Those dead on arrival (DOA) at the hospital had a BAC testing rate of 99 per cent, and those dead at the scene (DAS) of injury and taken directly to the county morgue had a BAC testing rate of 93 per cent. Persons admitted to a hospital with severe, moderate, or mild injuries had BAC testing rates of 61, 54, or 30 per cent, respectively.

BAC testing rates also varied by external cause of the injury (Figure 1). Persons with brain injuries from firearms had a BAC testing rate of 87 per cent, compared with 47 per cent from motor-vehicle crashes, 43 per cent from assault, 33 per cent from falls, and 22 per cent for all other external causes combined. Since the comparison of these BAC testing rates may be confounded by age, gender, and severity of brain injury, the external cause-specific rates were adjusted (by the direct method) for these factors. As shown in Figure 1, the differences in BAC testing rates among injury types (defined by external cause) were reduced substantially after adjustment, especially for firearm injuries.

BAC testing rates were highest for brain-injured pedestrians and lowest for bicyclists struck by a motor vehicle (Table 1). BAC testing rates were higher for drivers than for passengers, and the rates increased with severity of brain injury for all road-user categories. These BAC testing rates did not change appreciably when adjusted for age and gender.

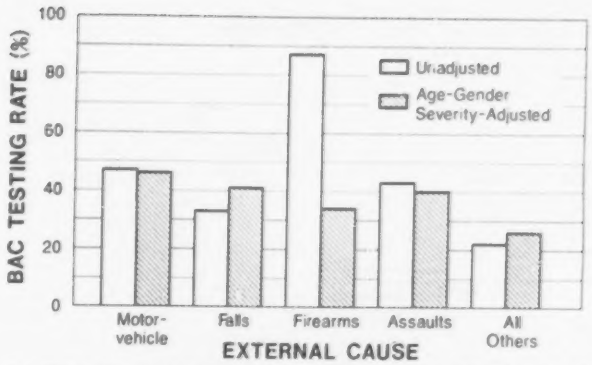

FIGURE 1—The per cent of brain-injured persons tested for BAC, by external cause of the injury: A comparison of results unadjusted and adjusted for age, gender, and injury severity

Proportion with Positive BAC

Of the 1,155 BAC-tested cases, 493 (43 per cent) had a negative result, 98 (8.5 per cent) had a BAC of 1 to 99 mg% (.22–21.70 mmol/L), 257 (22 percent) had a BAC of 100–199 mg% (21.71–43.20 mmol/L), and 307 (27 per cent) had a BAC of 200 mg% (43.42 mmol/L) or higher. Driving with BAC of 100 mg% (21.71 mmol/L) or higher is either an offense per se or presumptive evidence of driving under the influence (DUI). Driving with lower BACs can also result in DUI charges in most states. Generally, the proportion of those tested who had positive tests decreased beginning at about age 45 (Table 2). One-third of those aged 25–64 who tested BAC positive had a blood-alcohol level of 200 mg% (43.42 mmol/L) or higher.

Higher proportions of mildly brain-injured persons who were tested had positive BAC tests, compared with those with moderate or severe brain injuries or those who were DOA at the emergency room or taken directly to the coroner's office (Table 3).

The proportions with a positive blood-alcohol result varied by external cause of the injury. Two-thirds of those tested who were brain-injured from motor vehicle-related causes had a positive BAC; lower prevalences were observed for assaults (60 per cent), falls (44 per cent), firearms (35 per cent), and all other causes (42 per cent). These differences may be confounded by severity of injury and age (Tables 1 and 2); hence, the overall proportion with a positive BAC and the proportions at each BAC level were adjusted for these

TABLE 1—Per Cent of Brain-Injured Subjects Who Were Tested for Blood Alcohol Concentration (BAC), by Injury Severity and Road-User Category, San Diego County, California, 1981

Road-User Category	Injury Severity			
	Mild	Moderate/ Severe	DOA*	Total
Occupant	33	63	99	46
Driver	36	68	100	50
Passenger	29	56	96	39
Motorcyclist	34	66	100	49
Pedestrian	32	55	97	53
Bicyclist	20	29	100	29
Total	33	62	99	47

*Dead on arrival at the hospital.

TABLE 2—Per Cent of Tested Subjects with a Positive BAC and the Per Cent at Each BAC Level, by Age, San Diego County, California, 1981

	Blood Alcohol Concentration in mg% (mmol/L)			
Age (No. Tested)	1–99 (.2171–21.70)	100–199 (21.71–43.20)	200+ (43.21+)	Positive BAC* (≥1 mg% [.2171])
15–24 years (485)	11	26	22	58
25–44 (417)	8	26	33	67
45–64 (157)	8	12	34	54
65+ (96)	3	9	10	23
Total (1155)	9	22	27	58

*May not equal the sum of the other three columns because of rounding error.

factors. The unadjusted results did not differ much from the adjusted results for persons injured in motor vehicle crashes or falls (Figure 2) but did differ from the adjusted estimates for persons injured from firearms or assaults. Among tested subjects with a positive BAC, those injured by firearms tended to have a lower BAC level and those injured by falls had a higher BAC level than did other injured subjects.

Results of BAC testing by road-user category among motor vehicle-related brain injuries are given in Figure 3. The proportion with a positive BAC is between 60 and 70 per cent for all road-user categories, and the proportions are five to nine times higher for persons with BAC levels of 100 mg% (21.71 mmol/L) or higher than for those with BAC levels less than 100 mg% (21.71 mmol/L).

BAC results for brain-injured victims of motor vehicle crashes are given in Table 4 by injury severity and road-user category. Fifty-nine per cent of these tested subjects had BAC levels greater than or equal to 100 mg% (21.71 mmol/L), and this proportion was higher for drivers of cars, vans, or light trucks than for brain-injured passengers in these vehicles or for motorcyclists, pedestrians, and bicyclists. In each road-user category, the proportion having a high BAC level (≥100 mg% [21.71 mmol/L]) was greater for those with mild brain injuries than for those having more severe injuries.

Work-Related Brain Injuries

One hundred fourteen of the 2,649 members of the eligible cohort sustained a brain injury while engaged in work-related activities. Of the 21 subjects who were BAC-tested, 17 had a negative result, but all four subjects with a positive result had a BAC level of 100 mg% (21.71 mmol/L) or higher.

TABLE 3—Per Cent of Tested Subjects with a Positive BAC and the Per Cent at Each Level, by Injury Severity, San Diego County, California, 1981

	Blood Alcohol Concentration in mg% (mmol/L)			
Severity of Brain Injury	1–99 (.2171–21.70)	100–199 (21.71–43.20)	200+ (43.21+)	Positive BAC (≥1 mg% [.2171])
Mild	7	27	37	71
Moderate or Severe	8	18	21	48
DOA/DAS*	11	19	19	49
Total	9	22	27	58

*DOA = Dead on arrival at hospital; DAS = Dead at the scene of the injury.

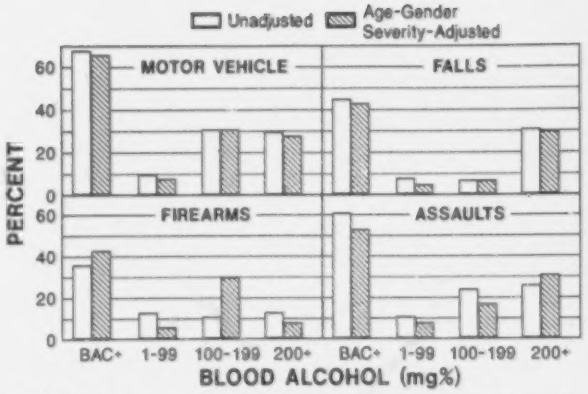

FIGURE 2—The per cent of brain-injured subjects with a positive BAC and the per cent of each BAC level, by external cause of the injury: A comparison of results unadjusted and adjusted for age, gender, and injury severity

BAC Testing and Possible Selection Bias

The effects of five factors on brain-injury severity (moderate or severe versus mild) are presented in Table 5. Results of multiple logistic regression show that the probability of sustaining a moderate or severe injury is positively associated with firearm injuries and age and inversely related to blood-alcohol level. This latter finding is consistent with the results of the crude (unadjusted) analyses shown in Tables 3 and 4. Although this inverse association with BAC level is in the direction opposite to that which one might expect, this finding may have resulted from differential rates of BAC testing by injury severity. It is conceivable, for example, that persons having a mild brain injury might not be tested for blood alcohol unless there was obvious intoxication while most persons with more severe injuries get routinely tested, regardless of actual BAC level. This explanation of selection bias is consistent with our data, which showed that while only 30 per cent of mild brain-injured persons were tested, 70 per cent of these persons had a positive BAC; yet, of the 74 per cent of moderate or severe injuries that were tested, only 45 per cent of these had a positive BAC. Thus, the apparent protective effect of alcohol seen in Table 5 (risk odds ratio of 0.6) could be the result of differential testing rates by severity.

The potential impact of differential BAC testing by severity is illustrated by the hypothetical data in Table 6. As shown in the upper panel, we have assumed that while there is no association between actual BAC level and BAC testing among all persons with moderate and severe injuries in the total source population (OR = 1.00), there is a positive association among persons with mild injuries (OR = 4.51). Using only the data for BAC-tested persons, as shown in the lower-left panel, we find that a positive BAC appears to protect against a more severe injury (OR = 0.35), which is the crude result actually obtained in this study. Using the hypothetical data for all persons in the source population, however, we see in the lower-right panel that there is no association between BAC level and injury severity (OR = 1.00). Thus, the inverse association observed in this study can be explained by selection bias caused by differential rates of BAC testing. Although the illustration in Table 6 is based on certain hypothetical data, we believe the underlying assumptions about BAC testing are quite plausible.

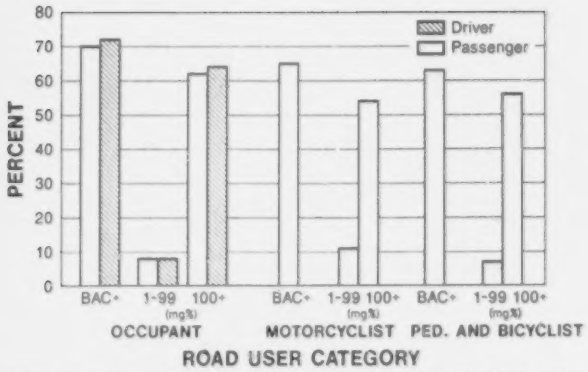

FIGURE 3—The per cent of motor vehicle-related brain-injured subjects with a positive BAC and the per cent at each BAC level, by road-user category and driver/passenger status

Effects on Hospital Mortality

Of the 1,155 persons who were tested for BAC level, 807 were admitted to a hospital alive because of their injuries and 86 died while hospitalized, an overall case-fatality rate of 11 per cent. The estimated effects of four factors on hospital mortality, based on multiple logistic regression, are given in Table 7. While the effects of injury severity and age are positive, as expected, the apparent protective effect of a higher BAC level on case fatality is similar to its effect on injury severity (Table 5). Thus, as with the previous finding, the observed effect of BAC level on case fatality might have been biased by differential rates of BAC testing (Table 6).

In Table 8, the adjusted BAC effect on hospital mortality is compared for various restrictions of the study population according to injury severity. Note that the apparent effect

TABLE 4—Per Cent of Tested Subjects with a Positive BAC and the Per Cent of Tested Subjects with a High BAC Level (≥100 mg% or 21.71+ mmol/L) among Brain-Injured Victims of Motor Vehicle Crashes, by Injury Severity and Road-User Category, San Diego County, California, 1981

Road-User Category	Severity of Brain Injury			
	Mild	Moderate/ Severe	DOA[a]	Total
Motor Vehicle Occupants[b]				
Positive BAC	80	60	59	70
≥100 mg%	73	51	48	62
Drivers[b]				
Positive BAC	82	60	63	72
≥100 mg%	76	52	52	64
Passengers[b]				
Positive BAC	74	60	48	65
≥100 mg%	66	50	33*	55
Motorcyclists[c]				
Positive BAC	72	52	74	66
≥100 mg%	63	43	58	54
Pedestrians and Bicyclists[d]				
Positive BAC	82	29*	59	68
≥100 mg%	74	29*	47	53
All Road Users				
Positive BAC	78	54	62	67
≥100 mg%	71	46	50	59

[a]Includes those dead on arrival at the hospital or taken directly to morgue.
[b]Includes passenger cars, vans, pickup trucks.
[c]Includes three passengers.
[d]Includes only those involved in a motor-vehicle collision.
*Based on fewer than 10 BAC-positive cases.

TABLE 5—Estimated Adjusted Effects of Five Factors on Injury Severity (Moderate/Severe vs Mild): Results of Multiple Logistic Regression (n = 1155)

Factor	Risk Odds Ratio (ROR)	95% CI (ROR)
Age	1.54*	(1.21, 1.96)
BAC Level (mg%)	0.60**	(0.53, 0.68)
Cause of Injury		
Firearms	9.92	(3.69, 26.7)
Falls	0.67	(0.54, 0.83)
Assaults	0.44	(0.36, 0.55)
Other***	1.00	

*Corresponds to an age difference of 25 years.
**Corresponds to a BAC level difference of 100 mg%.
***The reference category of which 78% are motor vehicle injuries.

(ROR <1) decreases (i.e., ROR becomes closer to one) as the logistic regression analysis is restricted to more severe injuries. This finding is also consistent with our previous interpretation of possible selection bias due to differential rates of BAC testing since most severe injuries are routinely tested.

Other Outcomes

Physician-diagnosed neurological impairments were recorded for all brain-injured patients admitted to a hospital with moderate or severe injuries *and* discharged alive from an acute care hospital. Rates of neurological limitations were higher among those patients with positive BAC levels (46 per cent) than for patients with negative BAC levels (32 per cent). The difference in rates was 14 per cent (95% CI = 2, 26). As shown in Figure 4, however, the elevated risk of neurological impairment (adjusted for age, gender, and severity) seems to be restricted to persons with a BAC level of 100 mg% (21.71 mmol/L) or greater.

TABLE 6—Hypothetical Illustration of Selection Bias Due to Differential BAC Testing in Estimating the Effect of BAC Level on Injury Severity

A. BAC Testing

Actual BAC Level	Moderate/Severe Injury			Mild Injury		
	BAC Test	No Test	Total	BAC Test	No Test	Total
Positive	275	95*	370*	384	444*	828*
Negative	332	115*	447*	161	840*	1001*
Total	607	210	817	545	1284	1829
OR = 1.00				OR = 4.51		

B. Brain Injury Severity

BAC Level	Observed among Tested Subjects			Hypothetical Data in Total Source Population (Tested and Untested)		
	Moderate/ Severe	Mild	Total	Moderate/ Severe	Mild	Total
Positive	275	384	659	370	828	1198
Negative	332	161	493	447	1001	1448
Total	607	545	1152	817	1829	2646
OR = 0.35 (95% CI = 0.27, 0.44)				OR = 1.00		

*These numbers are hypothetical but consistent with the assumptions about BAC testing that are stated in the text.

TABLE 7—Estimated Adjusted Effects of Five Predictors on Hospital Mortality among Hospitalized Patients: Results of Multiple Logistic Regression (n = 808)

Predictor	Risk Odds Ratio (ROR)	95% CI (ROR)
Age	2.90*	(1.70, 4.93)
Sex (Male/Female)	1.25	(0.85, 1.83)
BAC Level (mg%)	0.67**	(0.49, 0.92)
Injury Severity		
Severe	28.72	(10.4, 79.1)
Moderate	4.26	(1.46, 12.4)
Mild***	1.00	

*Corresponds to an age difference of 25 years.
**Corresponds to a BAC level difference of 100 mg%.
***The reference category.

Length of hospital stay did not differ by BAC level on admission. A similar result was found when persons who died in the hospital were excluded from the analysis. Yet length of stay was positively associated with BAC level among patients with severe injuries. Almost 80 per cent of those with a positive BAC test were hospitalized two weeks or longer, compared to 62 per cent of those with no blood alcohol on hospital admission. There was little association between length of stay and BAC level among patients with mild or moderate brain injuries.

Although the effect of BAC level on length of stay was observed for patients with severe brain injuries (i.e., GCS <8), it is still possible that this effect might have been confounded by amount or type of brain damage. Nevertheless, we found no evidence for such residual confounding in our data since the mean Injury Severity Score (ISS) in patients with severe injuries was similar for patients with positive and negative BAC levels.

Discussion

Findings from this study show that over one-half of all brain-injured persons who are seen in a hospital emergency room are BAC-tested; but until BAC testing for all hospital-treated injured persons is instituted, unbiased estimates of actual BAC frequency and its effects among brain-injured persons will not be possible.

We found that the rate of blood alcohol testing among brain-injured persons in San Diego County is related to type and severity of the injury. Thus, the inverse association observed between BAC level and injury outcome (severity and mortality) is very likely biased due to differential rates of BAC testing, especially among persons with less severe injuries (Table 6). Indeed, the adjusted estimate of the BAC effect on hospital mortality is reduced appreciably when the analysis is restricted to severe injuries only (Table 8).

TABLE 8—Comparison of Adjusted BAC Effect on Hospital Mortality by Restriction of the Study Population According to Injury Severity: Results of Three Multiple Logistic Regressions*

Severity Levels	No. Subjects	No. Deaths	Risk Odds Ratio (ROR)**	95% CI (ROR)
Mild, Moderate and Severe	808	88	0.67	0.49, 0.92
Moderate and Severe	263	87	0.80	0.61, 1.05
Severe Only	141	80	0.82	0.62, 1.09

*Each BAC effect estimate is adjusted for age, sex, and severity level (where needed).
**Corresponds to a BAC level difference of 100 mg%.

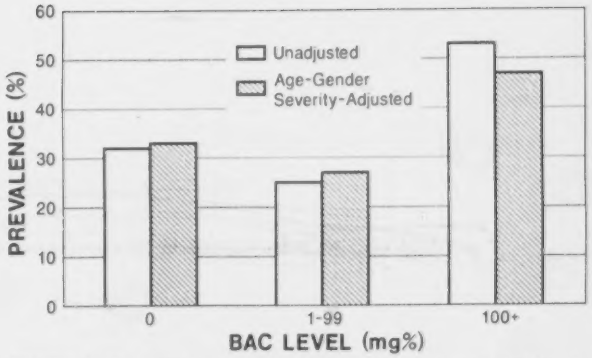

FIGURE 4—The prevalence (%) of physician-diagnosed neurological impairment at hospital discharge among moderately and severely brain-injured patients discharged alive from an acute care hospital, by BAC level: A comparison of results unadjusted and adjusted for age, gender, and injury severity

The ability of clinicians to diagnose and assess the nature and severity of brain injuries[35] influences treatment, particularly in persons with specific types of lesions.[16,23] Information on blood alcohol concentration is an important medical management tool and should be available on all persons seen for brain injuries in hospital emergency rooms.

There are at least two reasons why routine BAC testing of hospital emergency room patients who have been injured has not been more widely practiced:

- The principal function of emergency room activity is to determine whether the patient should be admitted to the hospital or discharged. If BAC testing is essential to the physician's decision on disposition of the patient, blood specimens are taken. Even if the patient shows signs of alcohol use (e.g., breath odor) the BAC test often may not be requested if it would not materially assist in determining disposition.
- Requesting a BAC test in the emergency room may have legal implications. That is, some physicians might not request a BAC test if the result, once recorded in the medical record, could lead to a legal problem for the patient. Because the prevailing practice in most emergency rooms of Level I or Level II trauma centers is to order a computerized tomography scan for any patient with altered consciousness, the value of a blood alcohol test is diminished. If there is altered consciousness and no overt evidence of alcohol use, however, a blood test may be requested to rule out alcohol in the diagnosis.

What remains for future research is to study the advantages and disadvantages of knowing blood alcohol level among possible brain-injured patients in the emergency room. Because blood samples are taken from patients with serious brain injuries who arrive at an emergency room for treatment of injuries, BAC testing could easily become routine even when there has been a substantial delay between injury and initiation of treatment. An alternative approach might be a quick, inexpensive, and noninvasive breath test to differentiate alcohol from ketones, as in diabetic ketosis.

The most perplexing finding from this study relates to outcomes from brain injury with and without alcohol involvement. Despite the animal evidence from observational and controlled laboratory studies,[5–15] the evidence that alcohol

affects injury outcome is not conclusive. Ward and colleagues[16] reported that in a large series of patients having many different types of injuries, mortality was lower in patients who had been drinking than in those who had not been drinking. Huth *et al*,[17] found that hospital course and late outcomes were unaffected by the patient's blood alcohol level. Luna and associates[18] found that although injury severity in critically head-injured motorcyclists was not related to alcohol intoxication, mortality was twice as high in intoxicated patients as in nonintoxicated patients. The investigators suggested that severe brain injury and acute ethanol intoxication acted in a synergistic manner to prevent the patient's homeostatic injury response. Results from these studies may also have been biased due to differential rates of BAC testing, as in this investigation.

Although injury severity and case fatality were inversely related to BAC level, the rate of neurological limitations among moderately or severely brain-injured survivors was positively related to BAC level. In addition, for those with severe brain injuries, controlling for age and total body injury severity, brain-injured subjects with positive blood-alcohol concentration had a longer duration of hospital stay. Since BAC testing rates were highest in the most seriously injured group, these observed effects are less likely to have resulted from selection bias. On the other hand, a physician's knowledge of alcohol use (or blood alcohol) at the time of injury might conceivably influence reporting of neurological limitations, which might be related to chronic alcohol use. But this alternative explanation is not very likely among younger patients with brain injuries.

It was not possible with the information available to identify chronic alcoholics from occasional drinkers among the brain-injured cohort. Hervé, *et al*,[20] reported recently on injury mortality and blood alcohol levels in France. Their data seem to suggest that chronic alcoholics had a higher mortality rate than did non-alcoholics, occasional drinkers, or non-drinkers, controlling for age and overall injury severity level. The possible interaction of long-term, heavy alcohol use and the acute intoxication at the time of an injury has not yet been investigated.

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Risk of Fatal Rollover in Utility Vehicles Relative to Static Stability

LEON S. ROBERTSON, PHD

Abstract: The risk of fatal rollover of utility vehicles per 100,000 registered vehicles relative to cars during 1982–87 was strongly correlated to the static stability of the vehicles. Distance between the center of the tires divided by twice the height of center of gravity explained 62 per cent of the variation in fatal rollover rates where rollover was the first harmful event. Statistical controls for 20 major risk factors indicated no correlations that would deflate the correlation between stability and rollover. Low stability utility vehicles roll over more often on the road suggesting that the lateral force of turning is often the tipping force. (*Am J Public Health* 1989; 79:300–303.)

Introduction

The sales of utility vehicles, sometimes called multipurpose vehicles, have increased substantially in recent years. In 1986 some 725,000 were sold in the United States compared to 132,000 in 1982.[1] Questions regarding the stability of these vehicles were raised in 1980 when the Jeep CJ-5, the Ford Bronco, and the Chevrolet Blazer were found to have rollover rates per registered vehicle much higher than passenger cars. These utility vehicles also had lower static stability than a sample of cars, as measured by the distance between the center of the tires divided by twice the height of center of gravity, usually expressed as T/2H.[2] T/2H of utility vehicles range from 1 to 1.2 while that of the vast majority of cars is above 1.2 and range as high as 1.6.[2,3]

The principle of static stability can be illustrated easily with a section of 2×4 lumber. Placed on its two inch side, it is easily tipped over by a force from the side, but placed on its four inch side, it tends to slide rather than tip over given the same force. The higher the weight from the ground relative to the width of the bottom decreases the stability. In a moving vehicle, the lateral force generated in a turn provides a potentially tipping force.

Studies of utility vehicles in the early 1980s confirmed the markedly higher rollover rates of utility vehicles relative to cars but did not attempt to correlate the rates with static stability.[4–6] A recent study found that, for the four utility vehicles and 11 cars for which static stability had been published, the fatal rollover rate per registered vehicles during 1981–84 in the US decreased exponentially in relation to static stability. The correlation was almost perfect; static stability explained 96 per cent of the variation among those vehicles. Static stability was not correlated to fatal crash rates where rollover did not occur.[7]

In response to a petition from Congress, the National Highway Traffic Safety Administration (NHTSA) gathered static stability data on a number of vehicles and conducted a study of single-vehicle crashes, mostly nonfatal, in three states. Static stability explained 86 per cent of the variation in rollover per cent.[3]

Since the relative risk of fatal rollover crashes in a number of recently marketed utility vehicles has not been studied, this communication reports an analysis of fatal rollover crashes in relation to static stability of these vehicles, as well as the older ones, during 1982–87 in the US. Other known risk factors were examined as possible confounding factors.

Method

The available data on static stability of utility vehicles that have been measured are presented in Table 1. In some cases, only one vehicle was measured. In the instances of multiple measures, the major intravehicle variation occurs in the vehicles that have the same name for different sized versions, namely Blazers and Broncos. It is possible that some of these vehicles were misidentified at time of measurement or in transmitting the information to NHTSA. In the analysis here, the median value rather than an average is used to minimize the effect of atypical measures or erroneous vehicle identification.

The data on fatal crashes of the utility vehicles in which an occupant of the vehicle died were extracted from the Fatal Accident Reporting System (FARS) data tapes for the calendar years 1982–87.[8] These tapes contain information on virtually every fatal crash in the US with the possible exception of some that occurred in late 1987 and had not been reported when the tape for that year was released. Vehicles were identified by decoding vehicle identification numbers that contain codes specifying makes and models. For comparison, all crashes in which an occupant died in rollovers and nonrollovers were also counted.

Data gathered from state motor vehicle administrations by R.L. Polk Company were used to count registrations by make and model during 1982–84.[9] Since R.L. Polk data do not include vehicles registered in Oklahoma, vehicles that crashed in Oklahoma during the years that registrations were counted from that source were eliminated from the analysis.

TABLE 1—High, Low and Median Stability Ratios (T/2H) Measured by Various Sources

Vehicle	High	Year	Source	Low	Year	Source	Median
1. CJ-5	1.01	79	Snyder	0.99	82	Ford	1.00
2. CJ-7	1.10	81	Unk	1.07	79	Snyder	1.08
3. Pre-78 Bronco	1.10	73	Snyder	1.07	71	Ford	1.08
4. Blazer/Jimmy	1.22	70	TRC	1.16	82	Unk.	1.19
5. Bronco II	1.12	84	Unk	1.01	83	STI	1.07
6. S-Blazer/Jimmy	1.20	85	TRC	1.04	?	Ford	1.07
7. Samurai	1.12	85	STI	1.12	85	STI	1.12
8. Montero	1.07	83	Unk	1.04	85	STI	1.05
9. Wrangler	1.13	87	TRC	1.13	87	TRC	1.13
10. 79+ Bronco	1.12	81	Unk	1.06	83	STI	1.09
11. Cherokee	1.19	83	TRC	1.19	83	TRC	1.19
12. Ramcharger	1.19	79	TRC	1.19	79	TRC	1.19
13. Trouper	1.10	87	TRC	1.10	87	TRC	1.10

DATA SOURCES: Snyder is reference 2; Ford = (Ford Motor Company); STI = (Systems Technology, Inc); TRC = (Transportation Research Center of Ohio); Unk = "unknown source" in the publicly available NHTSA files on rollover petitions.

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The regression coefficient of yearly changes in vehicle registrations by vehicle age was used to estimate scrappage of vehicles as they aged in 1985–87 (per cent change = −0.28 × vehicle age, $R^2 = 0.31$). For new vehicles, vehicle sales data[1,10] were used to count registrations. The monthly sales were multiplied by the number of months to the end of the study period and the sum of these months was divided by 12 to obtain registered years.

The rate of fatal rollover as "first harmful event," defined as the event causing the initial damage to the vehicle or occupants,[8] was examined for each vehicle with sufficient registrations relative to that of all cars. Log odds of the relative risk was used to establish confidence intervals.

If instability contributes mainly to rollover, there should be little increased risk of nonrollover fatal crash rates of vehicles with low stability ratios. Therefore, the relative risk of nonrollover fatal crashes of the utility vehicles relative to cars was examined as well.

Direct controls for the potential confounding effects of driver's characteristics or driving environment could not be applied since the necessary data on use were unavailable. Maximum potential confounding was estimated for 20 factors by measuring the relation between each factor and stability within the group of fatal rollover crashes. Factors unrelated to stability could not confound the effect of stability on risk. Expressed mathematically:

$$C\ L/H = RL/RH = b(S)$$

where L = low exposure to a risk factor
H = high exposure to a risk factor
C = constant ratio of risk from low to high risk situations
RL = fatal rollovers in low risk-factor situations
RH = fatal rollovers in high risk-factor situations
S = Stability value for a given vehicle
b = the slope of the correlation

This equation says that the ratio of rollovers in low relative to high-risk-factor situations is a function of the ratio of exposures in those situations times the relative risk ratio C. For other risk factors to explain the correlation of static stability to rollover rates, there must be a positive correlation of one or more of the ratios of rollovers by risk factor with static stability. Therefore, if b has a positive value, then the risk factor is a potential confounding factor that inflates the stability effect. A near-zero correlation indicates that the risk factor could not have produced a spurious correlation of stability and rollover rates. A negative correlation implies that the stability effect may be deflated by confounding. The ratio of numbers of rollovers in low relative to high-risk situations were examined for 20 major, known risk factors, including any alcohol and alcohol above the legal limit in the states where 80 per cent or more of fatally injured drivers were tested.

Results

Vehicles that had fewer than 100,000 registered years during the study period—the Montero, Trouper, and Wrangler—were eliminated from the analysis. Table 2 presents the analysis of the relative risk of fatal rollover as first harmful event in relation to the all-car rate, 1.3 per 100,000 registered vehicles. The Jeep CJ-5 was 19.7 times more likely to experience a fatal rollover per registered vehicle than a car. The Jeep CJ-7, the pre-78 Ford Bronco, and the Ford BroncoII were clustered in the range of 10 to 12 times the fatal rollover rate of cars. Another cluster with a relative risk three to six times that of cars included the Dodge Ramcharger, the 1982–87 Ford Bronco, the General Motors large and small Blazer/Jimmys and the Suzuki Samurai. The lowest relative risk was 1.3 for the Jeep Cherokee and the lower confidence limit for that vehicle was the only one within the range of all cars.

The relation of the first-harmful-event rollover rate to the stability ratio, T/2H, displayed in Figure 1, is striking. The higher the stability the lower the fatal rollover rate. In a linear model, 62 per cent of the variation among vehicles was accounted for by the stability ratio. The regression coefficient of −86 (95% confidence interval = −39 to −132) suggests that increasing the stability ratio by one-tenth decreases the annual fatal rollover rate by an average of about 9 per 100,000 registered vehicles, within the range of stability ratios of the vehicles in the study.

Since vehicles sometimes roll over subsequent to other "first harmful events," the correlation of stability ratios with rate of rollover as "most harmful event" was also examined. The pattern was similar to that in Figure 1 but the rates were higher.* The regression coefficient of −143 (95% CI = −64 to −222, $R^2 = 0.61$) suggests that an increase in the stability ratio of one-tenth would decrease the annual most-harmful-event rollover rate in which an occupant dies by an average 14 per 100,000 vehicles, again within the range of stability ratios of these vehicles.

The risk ratios for utility vehicle occupants relative to car occupants of fatal crashes without rollover are displayed in Table 3. Eight of the 10 vehicles had risk ratios less than one relative to cars, two with upper bound less than one. Only the Jeep CJs had confidence intervals above and outside the range of the all-car rate and their relative risk ratios were less than two.

The relation between stability and driver or environmental characteristics is shown in Table 4. None of the 20 risk factors was substantially positively correlated to T/2H. The on-road/left-road correlation was negative, indicating that lower stability vehicles more often rolled over on the road rather than after leaving the road, as plotted in Figure 2.

The other environmental risk factors were unrelated systematically to vehicle stability. No driver factor was correlated systematically to vehicle stability, including age, gender, presence of alcohol or illegal alcohol, invalid licenses, prior crashes, prior suspensions or convictions for driving while intoxicated, speeding, or other offenses.

*Data available on request to the author.

TABLE 2—Relative Risks of Fatal Rollover Crashes

Vehicle Model	Model Years	Fatal Rollovers	100,000 Year Registered	Relative Risk	95% CI
All passenger cars	all	9813	7790.6	1.0	
Jeep CJ-5	67–83	354	14.3	19.7	17.7–21.9
Jeep CJ-7	76–86	216	13.8	12.4	10.8–14.2
Ford Bronco	74–77	84	5.9	11.3	9.1–14.0
Ford Bronco II	83–87	141	11.0	10.2	8.6–12.1
Suzuki Samurai	86–87	9	1.2	6.1	3.2–11.8
Ford Bronco	82–87	51	7.1	5.7	4.3– 7.5
GM Blazer/Jimmy	69–87	222	32.8	5.4	4.7– 6.1
GM S-Blazer/ S-Jimmy	82–87	89	22.1	3.2	2.6– 3.9
Dodge Ramcharger	82–87	15	3.8	3.1	1.9– 5.2
Jeep Cherokee	75–83	11	6.9	1.3	0.7– 2.3

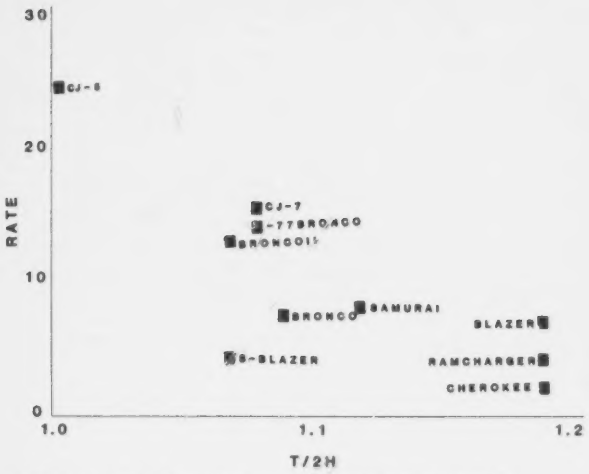

FIGURE 1—Fatal Rollover Crashes per 100,000 Registered Vehicles

Discussion

These and previously published data leave little doubt that vehicle stability is the primary factor contributing to the higher fatal rollover crashes in utility vehicles than cars. The result is consistent with well-known principles of physics. The nonrollover fatal crash rate of utility vehicles, which is caused by factors other than stability, is similar to that of cars. None of the powerful predictors of fatal crashes generally—rural roads,[11] lack of light,[12] prior driving record,[13] alcohol,[14] age and gender[15]—were related to stability ratios in such a way as to be confounding factors.

The only risk factor related to stability is the on-road/off-road factor. The fact that the less stable vehicles more often roll over on the road is consistent with the hypothesis that many more of them roll over because of the lateral force of turning alone, or perhaps even occasionally by very strong side winds, rather than going over embankments or ramping, which is much more likely to happen off the road rather than on it.

The possibility that some combination of risk factors contributes to the correlation of stability and rollover is extremely remote. The factors most strongly predictive of fatal crashes generally, such as age and gender, are known to be independently correlated to the risk of a fatal crash.[12] While all of the risk factors examined do not occur independently of one another, the fatal rollover crashes are distributed among the categories of the factors in such a way that high enough concentration in some important combination, which is also in turn strongly combined with low stability vehicles, is not possible given the lack of correlation of stability with the factors considered separately.

Despite widespread publicity in 1980 regarding the rollover propensity of the Jeep CJs, several manufacturers have subsequently introduced new vehicles with low stability. The lack of a perfect correlation of rollover rates with static stability suggests that some manufacturers of later model vehicles may have partially offset the effect with better than average suspension systems or other vehicle characteristics moderating the tendency of lower stability vehicles to

TABLE 3—Relative Risks of Fatal Nonrollover Crashes to Occupants of Utility Vehicles Relative to the All-Car Rate

Vehicle Model	Model Years	Fatal Nonrollovers	Relative Risk	95% CI
All passengers cars	all	100,886	1.0	
Jeep CJ-5	67–83	320	1.7	1.5–1.9
Jeep CJ-7	76–86	244	1.4	1.2–1.6
Ford Bronco	74–77	64	0.8	0.7–1.1
Ford Bronco II	83–87	96	0.7	0.6–0.8
Suzuki Samurai	86–87	14	0.9	0.6–1.6
Ford Bronco	82–87	77	0.8	0.7–1.0
GM Blazer/Jimmy	69–87	335	0.8	0.7–0.9
GM S-Blazer/ S-Jimmy	82–87	245	0.9	0.8–1.0
Dodge Ramcharger	82–87	35	0.7	0.5–1.0
Jeep Cherokee	75–83	77	0.8	0.7–1.0

TABLE 4—Regression Analysis of Ratio of Number of Rollovers in Low- and High-Risk Situations by Static Stability

Environment	T/2H Coefficient	95% CI Lower	95% CI Upper	% Variance Explained
Urban/Rural	−0.67	−1.89	0.56	14
Interstate/Other	1.91	−0.40	4.23	27
On Road/Off Road	−3.96	−7.27	−0.64	44
3+ lanes/2 lane	0.05	−0.30	0.40	1
Speed limit <55/55+	−1.39	−2.91	0.12	32
Straight/Curve	3.40	−1.73	8.54	19
Level/Grade	−1.46	−8.47	5.55	2
Concrete/Blacktop	0.10	−0.66	0.86	1
Dry/Wet	−14.68	−75.47	46.11	3
Daylight/Other	2.05	−0.89	4.99	21
Driver				
Valid License/Other	0.56	−0.36	1.47	17
No Prior Crash/1+	−8.52	−24.31	7.28	14
No Prior Suspension/1+	−29.49	−85.07	26.10	15
No Prior DWI/1+	−27.89	−71.07	15.29	21
No Prior Speeding/1+	8.78	−4.22	21.79	18
No Other Conviction/1+	6.30	−16.53	29.13	4
No Blood Alcohol/0.01+	−0.21	−3.49	3.07	3
No Illegal BAC/0.10+	1.23	−4.93	7.39	3
<25 years old/25+	6.41	−3.32	16.15	17
Women/Men	−0.27	−1.76	1.22	2

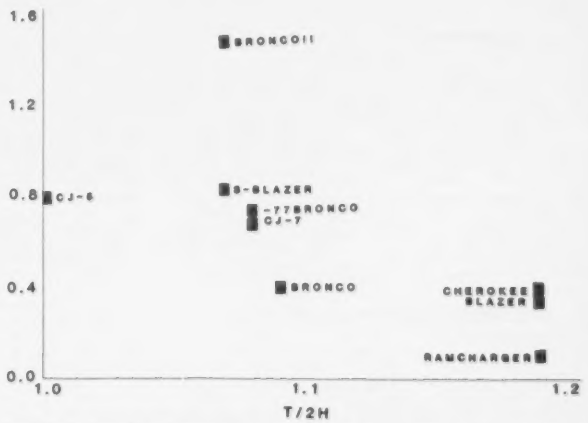

FIGURE 2—Ratio of Rollovers on the Road to Rollovers After Leaving the Road (Excludes the Samurai for which there were less than 10 cases)

tip over in turns. Nevertheless, the static stability is a strong predictor that could be used to set standards for vehicles.

The National Highway Traffic Safety Administration has denied petitions to address the problem.[16,17] The agency's main justification for its inaction was that static stability is a continuously distributed variable and that to choose a required point would be arbitrary. It also said that it is prohibited by law from banning a class of vehicles which such a standard would do. If standards could not be set on continuously distributed factors, however, there would be no standards for blood alcohol concentration and head injury criteria. Since a standard for static stability could be met by simply widening the distance between the center of the tires and/or lowering the center of gravity, such a standard would not ban a class of vehicles. Furthermore, one utility vehicle with stability in the range of that of passenger cars had a rollover rate similar to passenger cars.

Another issue is whether the vehicles now in use should be recalled and modified to reduce rollover risk. Smaller diameter tires would reduce the height of the center of gravity on the high-risk vehicles. It may also be possible to weld some weight along the undersides of the vehicles to make them more stable. Given the risk and the fact that each cohort of these vehicles will be in use an average of more than 10 years, a testing program seems indicated to evaluate the effect of such modifications on stability.

ACKNOWLEDGMENT

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4th National Environmental Health Conference, June 20–23

The Center for Environmental Health and Injury Control of the Centers for Disease Control; the Agency for Toxic Substances and Disease Registry; and the Association of State and Territorial Health Officials have announced the Fourth National Environmental Health Conference will be held June 20–23, 1989, in San Antonio, Texas. The primary audience is federal, state, and local health and environmental officials and physicians and the environmental community.

The theme of the 1989 conference is "Environmental Issues: Today's Challenge for the Future." The objectives are to: address the environmental insults that have the greatest importance to public health; review topical scientific findings; and discuss prevention strategies. Plenary sessions will be on radon; medical, municipal, and hazardous waste; air pollution; lead in the environment; and dioxin. Twenty workshops will be held on topics of interest to states, academic institutions, and federal agencies including health assessments at NPL (national priorities list) and RCRA (Resource Conservation and Recovery Act) sites, emergency responding, radiation, birth defects, risk communication, indoor air pollution, and respiratory disease.

Category 1 Continuing Medical Education credits for physicians only are being offered for designated sessions sponsored by the Centers for Disease Control, and accredited by the Accreditation Council for Continuing Medical Education.

For further information, contact either Lewis Webb (404/488-4700) or Dr. John Andrews (404/488-4682) or the Centers for Disease Control or Peter McCumiskey (404/488-4682) at ATSDR, all in Atlanta, GA.

Evaluation of a Community-Wide Incentive Program to Promote Safety Restraint Use

ROBERT D. FOSS, PHD

Abstract: A program was designed and evaluated to determine the feasibility of implementing an incentive campaign to promote safety restraint use, primarily among children up to age 13. During a five-month period, community residents were eligible to win monthly prizes and a grand prize by buckling up all automobile occupants when driving. Systematic monitoring of traffic and a telephone survey indicate that the program was minimally effective, reaching the attention of only about half the families with children in the community despite extensive advertising. Effects of the program on seat belt use closely paralleled airing of radio announcements, appear to have been confined primarily to children under age four, and tailed off somewhat after the initial response. (*Am J Public Health* 1989; 79:304–306.)

Introduction

It has now been a decade since the first state law was passed mandating the use of safety seats to protect children from injury or death in automobiles.[1] All 50 states and the District of Columbia now have enacted child safety restraint laws and many of the earliest laws have been strengthened. Several defects notwithstanding,[2] the laws appear to have been modestly successful in increasing the number of very young children who ride properly buckled up, and in reducing injuries and fatalities.[3,4] Yet according to ongoing, nationwide observation of child safety restraint use, only 27 per cent and 17 per cent of 3-year and 4-year-old children, respectively, (most of them born after child safety seat laws were enacted in their states) rode buckled up in 1985.[5]

A number of experimental studies have found that short-term educational interventions do not reliably affect whether parents will buckle up their children beyond the first few months of life,[6–9] if that.[10,11] Nor is there any substantial evidence that the availability of seat loaner/rental programs, by themselves, serves to substantially increase the number of young children who ride buckled up.[12]

It has been suggested that rewarding safety restraint use provides an alternative way to increase the use of safety restraints.[13,14] Incentive programs have been developed in a number of communities,[15–17] based on the notion that many individuals who can be induced to buckle up repeatedly over a period of time, by the offer of a chance for material reward, will continue to buckle up out of habit even after rewards cease. Several careful studies have documented impressive short-term increases in adult safety belt use as a result of such incentive programs, both in limited populations and on a community-wide basis.[15,18,19]

Although most incentive programs have focused on adults, it has been demonstrated that both elementary and nursery school children can be induced to buckle up as well by carefully designed incentive programs.[14,20,21] One of the primary advantages of incentive programs is that they can be extremely cost-effective when focused on an entire community, rather than a small, delimited population such as a workplace, school, or day care center.[13] Adopting a community focus, however, requires somewhat different tactics, which could conceivably render an incentive program far less successful.

A community-wide incentive program, with several crucial differences from small scale incentive programs, was designed and implemented. In contrast to earlier studies, this program: focused on a wide age range (birth through age 12); offered only a small chance of winning monthly prizes and a grand prize, rather than a substantial chance for small rewards; did not maintain high visibility monitoring of a fixed location, but instead involved unobtrusive, temporally and geographically diffuse monitoring for prize eligibility; and was conducted (except for the evaluation component) by a small group of volunteers. All these elements may reduce effects of an incentive program, but are necessities in community-wide programs, especially those that are not part of a well-financed research project. Perhaps the largest obstacle to implementation of a community-wide project is effectively bringing it to the attention of a substantial proportion of the public, a task that is much more easily managed in programs designed for geographically confined populations.

Method

The program, conducted from mid-September through mid-February in a university community with a nonstudent population of approximately 45,000, was designed to provide incentives for parents to buckle up both their children and themselves, as well as to influence school age children directly. The incentives were prizes in a monthly drawing of license numbers for cars whose occupants had been spotted buckled up at some time during the month. Specifically, any car containing at least one child judged to be under the age of 13 that was spotted by a member of the local Child Passenger Safety Council, in which all occupants were buckled up, was entered in a monthly lottery. Each time a car that met the criterion was spotted, its license number was recorded and entered, so the more often everyone buckled up, the greater the likelihood of the car's owner winning a prize. On the tenth day of each month a winner was randomly selected from the license numbers of those cars that had been recorded during the preceding 30 days. Unless they won a prize, members of the target population had no way of knowing whether their behavior had been detected.

The program was advertised by two weeks of radio advertisements (30-second spots, six times daily), a newspaper article and a supportive newspaper editorial as the contest began, in addition to approximately 80 large posters (42.5 cm × 57.5 cm) placed prominently throughout the community. In addition, letter-size flyers were distributed monthly to several thousand elementary school children announcing the prize for the month as well as the grand prize. On two occasions these distributions were accompanied by a

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letter to children emphasizing the importance of seat belt use, explaining how they might win, and encouraging them to pressure their parents to buckle up and to be sure siblings also were restrained. Flyers were also posted monthly next to the larger posters. Several hundred flyers also were mailed directly to parents of young children <5 years old who had participated in two earlier telephone interview studies concerning safety restraint use.

Monthly prizes ranged in value from $50 to $200. The grand prize, for which every car observed with everyone buckled up at any time during the five-month period was eligible, was a family vacation to a Florida theme park. All prizes were donated by local merchants. Some radio time was donated by local stations; the remainder of the ads were purchased with funds donated by community members to support efforts of the local Child Passenger Safety Council. Although a state law mandating proper restraint of children up to age five was in effect at the time of the study, it was not enforced, nor had it ever been enforced or advertised since its enactment over three years earlier.

Prior to the contest, throughout its duration, and for some time thereafter, data on safety restraint use by children in the community were systematically collected by paid, trained observers. None of these observers were involved in spotting cars for prize eligibility. This ongoing monitoring activity was conducted for research purposes tangentially related to the incentive program. Observations were normally made, weather permitting, from 10 am to noon on Saturday mornings and from 3–5 pm during the week. Darkness and frigid weather reduced observation periods to one hour from November through February. Several additional late morning, mid-week observation periods were scheduled during Christmas break, when a reasonable number of older children could be observed. Observers were stationed at several locations that were selected for community representativeness and maximum visibility of automobile occupants. Observers recorded the restraint use, estimated age, and seating position of all children in automobiles that passed their vantage points. Although most observers probably were aware of the contest, they were not told that they were evaluating its effects and were unlikely to have drawn that conclusion on their own, since observations had been conducted for about 10 months before the contest began.

Results

School age children were analyzed separately from children under age six.

Infants and Preschoolers

From a baseline of 27 per cent prior to the contest, restraint of children under age six in vehicles rose immediately and dramatically in response to the early publicity announcing the contest, reaching 38 per cent during September and 43 per cent during October (Table 1).

About one month after initial radio advertising of the contest was discontinued, restraint use began to decline. It settled back to the baseline rate briefly, about 16 weeks after the contest began, and 13 weeks after media reminders of the contest had stopped. In response to this decline, an additional two weeks of radio ads were aired a few weeks prior to the end of the contest. Restraint use again increased, to 33 per cent during the final week of the contest. During the five-month incentive period, restraint use averaged nearly 10 per cent above the pre-incentive period. During the seven-week period immediately following the contest, a consistent rate 5 per cent above the pre-incentive period was maintained.

Because restraint use varies widely by age among preschool age children,[22–24] effects of the incentive program were examined for separate age groups. As Figure 1 illustrates, children above the age of four appear to have been affected initially, but the effect drops off later. Because age was estimated, these results must be interpreted with caution. Nonetheless, the pattern of sharp decline in use, and difference in responsiveness to incentives between infants and older children is unmistakable.

School-Age Children

Restraint use by children ages six through 12 years riding in rear seats (i.e., use of lap belts) is difficult to measure reliably in moving vehicles. Consequently, data are presented only for front seat occupants and there are fewer observations than for infants and preschoolers. The pre-incentive use rate (10 per cent) increased to 16 per cent during the second month of the contest, but declined again. During the final month of the contest, in concert with renewed radio advertising, the use rate peaked at nearly 20 per cent, declining again to 13 per cent during the seven weeks following the contest.

Although the incentive campaign required adults as well as children to be buckled up to qualify for prizes, safety belt use among adult front seat occupants hovered around 11 per cent before, during, and following the contest period.

Telephone Survey Results

To determine awareness of the incentive program, telephone interviews were conducted shortly after the campaign with 40 randomly selected parents of children from four of the schools where flyers had been distributed. Despite extensive advertising, only 59 per cent of the respondents were aware of the contest. Among those parents who did know of the incentive program, 35 per cent reported learning about it from the radio spots. The next most effective method was direct distribution of flyers to school children, with 26 per cent of

TABLE 1—Proportion of Children Restrained Before, During, and After Incentive Period by Age

	0–5 years (N)	Difference from Baseline (95% CI)	6–12 years (N)	Difference from Baseline (95% CI)
4 months Pre-incentive (Baseline)	.268 (921/3436)	—	.104 (104/1002)	—
September	.382 (164/429)	.114 (.069,.159)	.101 (21/208)	−.003 (−.048,.043)
October	.426 (266/624)	.158 (.119,.197)	.156 (62/397)	.052 (.015,.090)
November	.350 (165/471)	.082 (.039,.125)	.068 (7/103)	−.036 (−.097,.025)
December	.327 (337/1031)	.059 (.028,.090)	.082 (25/305)	−.022 (−.060,.016)
January	.266 (90/338)	−.002 (−.051,.048)	.197 (13/66)	.093 (.015,.171)
7 weeks Post-incentive	.316 (294/931)	.048 (.015,.080)	.126 (30/239)	.022 (−.02,.066)
Full incentive period	.364 (959/2636)	.096 (.072,.119)	.119 (128/1079)	.015 (−.012,.042)

NOTE: N given in parentheses. Contest began September 10, ended February 10; observations for September cover last three weeks of the month; observations for first week of February are included with January.

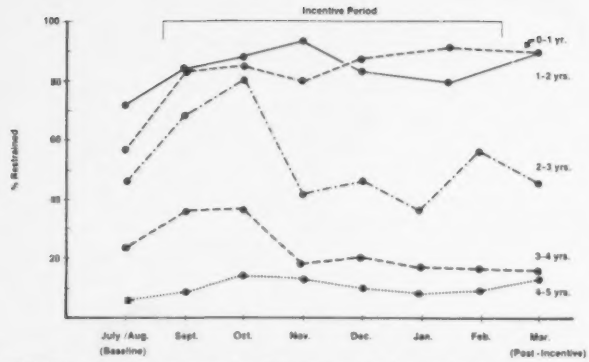

FIGURE 1—Effect of Incentive Program on Safety Restraint Use among Children under Five Years by Age. (Data were combined for January and February since an insufficient number of cases were observed for children below age two to provide stable estimates of monthly use within these age groups)

parents reporting that they learned about the contest from the flyers. Although 91 per cent of those who knew about the contest knew they needed to buckle up to win, only 18 per cent accurately reported the specific but crucial requirement that everyone in the car had to be buckled up. Almost nobody could identify the month the contest began, the month it ended, or any of the monthly prizes.

Discussion

The alterations needed to direct an incentive campaign to an entire community appear to attenuate the effectiveness of this approach. In the current case, creating and maintaining awareness of the program was a major problem. During periods of extensive radio advertising, response to the program was clear, although unspectacular. The fact that few people could identify information that was clearly printed on thousands of flyers distributed in the community and advertised repeatedly further emphasizes the difficulty of implementing an incentive campaign in a general population. At the least, it suggests that requirements must be extremely simple.

It appears that radio ads aired over a short period of time in a small community were an effective way to advertise the program. The modest resurgence in response to renewed ads suggests that it may be possible to sustain a moderate, community-wide increase in the rate of safety restraint use long enough for habitual use to be established. It remains to be seen whether this finding is generalizable to a more populated area. Th question of whether heavy advertising over a several month period may lose its effectiveness must also be addressed.

The marked differences in response to incentives to buckle up by various age groups point once again to the second, third, and fourth years of life as a pivotal time in safety restraint use.[13] These children, who are least likely to be restrained, are also the most difficult to reach with a programmatic effort. A similar age-differentiated effect was found for the state child restraint law a few years earlier in this same community.*

The third and fourth years of life are a time when children are in transition from near total control by parents to a point where they engage in many behaviors, including safety seat/belt use, substantially in accordance with only their own wishes. Although the body of literature on child development is huge, we know virtually nothing about how or whether this age-related increase in the autonomy of children is implicated in the distressingly high prevalence of nonuse of safety restraints by older preschool aged children. Nor has our understanding of parent-child interaction or of effective child-rearing been tapped. Such matters clearly must be understood if we are to deal effectively with safety belt use among children. Instead of merely demanding that parents protect their children, through laws mandating safety belt use, the solution to this problem probably lies in the development of techniques to encourage and assist parents to develop the habit of safety belt use by their children at a very early age.

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Evaluation of the Efficacy of Simulation Games in Traffic Safety Education of Kindergarten Children

LISE RENAUD, PHD, AND SAMY SUISSA, PHD

Abstract: Using a simulation game designed to teach children to obey certain traffic safety rules, an experimental study was conducted with 136 five-year-old children in four Quebec schools. Within each classroom, subjects were randomly divided into four groups: three intervention groups and one control group. Each of the experimental groups was subjected to a different intervention with outcome measured using three instruments related to attitudes, behavior, and transfer of learning of pedestrian traffic safety. Results suggest that simulation games including role-playing/group dynamics and modeling/training can change attitudes and modify behavior in the area of pedestrian traffic safety in children of this age. (*Am J Public Health* 1989; 79:307–309.)

Introduction

In Quebec, a Canadian province with 6 million inhabitants, the 1984 pedestrian injury rate in 5- to 9-year old children was 162.4/100,000 as compared to 67.1 for the population as a whole.[1] The prevention of such accidents involves legislation, engineering, education, or a combination of these. One possible way of preventing accidents is to teach children to obey certain rules of safety. This study evaluates a simulation game as an effective educational strategy in reaching this goal.

A "simulation game" is a simplified and dynamic model of a real or hypothetical system where players, following a set of rules, compete or cooperate in order to win. We used a simulation game, developed by la Régie de l'Assurance automobile du Québec (Automobile Insurance Board) and the Ministère de l'Education, and a proposed component of Quebec's preschool program.

Research on attitude and behavior modification suggests that role-playing group dynamics and model-training are effective means of triggering change in attitudes or behaviors and that they can be used effectively in simulation games. Role-playing group dynamics is more important in attitude change,[2–7] while modeling-training appears to have a stronger impact on behavior.[8–12] Simulation is a good strategy for effecting changes of attitudes or behaviors in psychosocial situations and produces transfer of learning.[13–16] Transfer of learning is defined as: "The extent to which the learning of an instructional event contributes or detracts from subsequent problem solving or learning of subsequent instructional events."[17] For example, knowledge in one sport may be transferable to another sport, typewriting skills may be transferred to a computer keyboard, etc.

The objective of this study was to determine the effect of certain elements within a simulation—namely role playing/group dynamics, and behavior modeling/training—on the attitudes and behaviors and transfer of learning of five-year-old pedestrians with respect to traffic safety rules.

Methods

We used an experimental scheme based on the post-test-only control group design because a pre-test might have provided participants with knowledge which would influence their reaction to the instrument.

The four schools participating in this study are representative of the schools in Montreal, a principal city in the province of Quebec. Within each school, two kindergarten classes composed of five-year-old children were selected. This age group was chosen because the traffic safety program is designed for kindergarten children. Within each class, children were randomly allocated to each of four groups: three intervention and one control.

The experiment consisted of two three-hour sessions. The first involved the presentation of the simulation games and the administration of attitude and behavior tests. The second session, conducted 10 days later, included only the administration of the transfer of learning test.

Intervention

We used three simulation games with the same learning content: traffic safety for the pedestrian. Each of the three experimental interventions included different elements within the same basic simulation game structure. Basically, the game simulates a part of a town complete with streets and sidewalks (dimensions: 1.22m × 2.44m). Five-year-old pedestrians "walk through" the miniature streets using finger puppets drawn on the backs of their hands.* The three simulation games are:

- *ATTITUDE:* This simulation game includes role-playing and group dynamics elements and is intended to primarily affect attitude change. Personal (verbal and behavioral) interactions and a loose scenario centered around a specific situation constitute the unique features of this simulation game.
- *BEHAVIOR:* This simulation game includes modeling and training elements and is intended to change behavior. Observation of a model, behavioral imitation, and feedback constitute the unique features of this simulation game.
- *ATTITUDE AND BEHAVIOR:* This simulation game includes role-playing, group dynamics, modeling and training elements and is intended to alter both attitude and behavior. The features from the other two simulation games constitute the simulation game.

Measurements

Three measurement instruments were used.* The instruments were simple since the experiment involved five-year-olds.

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*More details on the simulation game and measurement tests are available from the authors.

Attitude measurements were based on methods developed by Gochman.[18] Specifically designed for children, they measure the behavioral component of attitudes, i.e., the individual's predisposition to act when faced with the object of an attitude. This predisposition is determined by the child's perception of pedestrian vulnerability to traffic accidents, and the child's perception of the degree of danger that traffic in general presents to the pedestrian. Photographs are used to identify health risks. The children's perception of risk (health dimension) and their recommendations to avoid risk (prevention dimension) are measured. Ten photographs, validated by a jury of three researchers and tested in another group of children, were used to measure the perception of risk.

Behavior measurements were based on methods developed by Whalen and Rolf.[19] Children use images to show how they would act in a given situation. The measurement instrument is composed of a picture of a road, a series of questions, and stickers to be used in answering the questions.

Transfer of learning was measured using the methods developed by Haynes.[20] The transfer of learning was measured by observing children's reaction to a quasi-real life model of traffic risks set-up in the gymnasium. Only the lateral transfer of learning, which is defined as any facilitation of learning resulting from possession of previous knowledge in a related field, was studied. A trained observer notes the presence or absence of a predetermined selected set of behaviors using a scoring sheet specifically designed for the behavior being observed.

Analysis

The reliability of each measurement instrument was evaluated using Chronbach's alpha. Pearson's correlation coefficient was used to associate the two different dimensions of the attitude scale, as well as to verify that attitude, behavior and transfer of learning were not related.

The four intervention groups were compared, for each of attitude, behavior, and transfer of learning, using a three-way analysis of variance model (main factor: intervention; co-factors: school, sex). Since school was considered a random factor, a mixed model was used. Interactions and co-factors were tested and dropped from the final analysis if deemed insignificant. Multiple comparisons of confidence intervals were obtained using Tukey's method.

Results

The attitude measure had a range of 0–3. Each of its two dimensions, health and prevention, had a reliability coefficient of $\alpha = 0.89$ indicating good reliability. These dimensions were highly associated (Pearson correlation coefficient, $r = 0.61$), suggesting that they constitute two aspects of the same construct. The reliability coefficient for the global score of the behavior test (range: 0 to 5) is $\alpha = 0.41$ (the small number of items (5) in this test substantially affects the reliability coefficient). The reliability coefficient for the transfer of learning test (range: 0 to 31) is $\alpha = 0.85$.

Little association was found between the two dimensions of attitude and the other study variables (behavior and transfer of learning) (correlation of health dimension of attitude and behavior, $r = 0.20$; prevention dimension of attitude and behavior, $r = 0.23$, health dimension of attitude and transfer of learning, $r = 0.30$; and prevention dimension of attitude and transfer of learning, $r = 0.26$).

With respect to the primary hypotheses under study, the statistical assessment is made separately for each of the three components of outcome. In each case, sex was not significant and was removed from the analysis. In addition, the interaction between the intervention and the school factors was negligible and therefore removed from the analysis.

TABLE 1—Summary Statistics of Attitude, Behavior and Transfer Scores by Intervention Group

Variables	Intervention			
	Control	Attitude Simulation Game	Behavior Simulation Game	Attitude/ Behavior Simulation Game
ATTITUDE (range 0–3)				
Sample size	33	35	33	34
Mean ± SD	1.3 ± 0.7	1.9 ± 0.7	1.8 ± 0.6	2.0 ± 0.7
BEHAVIOR (range 0–5)				
Sample size	34	35	33	34
Mean ± SD	4.0 ± 1.1	4.9 ± 0.2	4.3 ± 0.8	4.5 ± 0.7
TRANSFER (range 0–31)				
Sample size	30	33	32	33
Mean ± SD	7.9 ± 3.7	8.7 ± 3.1	10.4 ± 2.1	10.1 ± 2.3

NOTES: Attitude simulation game = simulation game including role-playing/group dynamics.
Behavior simulation game = simulation game including modeling/training.
Attitude/behavior simulation game = simulation game including role-playing/group dynamics, and modeling/training.
Control = no intervention.

Attitude

Table 1 indicates that the means of the attitude test are similar between the intervention groups but quite different from those of the control group. Table 2 shows a clear difference between each of the three intervention groups and the control group on the attitude test. The group which used the simulation game involving role-playing/group dynamics, and modeling/training performed only slightly better than the two other intervention groups.

Behavior

The control group mean is lower than the three intervention group means. The means of groups using the simulation game that included attitude triggering elements (attitude and behavior simulation and attitude simulation) are

TABLE 2—Pairwise Mean Differences of Attitude, Behavior and Transfer Scores among Interventions for Pre-selected Contrasts

Variables	Difference Between Mean Scores	95% Confidence Interval
ATTITUDE		
Attitude simulation game vs control	0.60	0.13, 1.07
Attitude/behavior simulation game vs control	0.72	0.26, 1.21
Behavior simulation game vs control	0.52	0.05, 1.00
BEHAVIOR		
Behavior simulation game vs control	0.33	−0.57, 1.24
Attitude/behavior simulation game vs control	0.50	−0.40, 1.40
Attitude simulation game vs control	0.94	0.05, 1.84
TRANSFER		
Attitude simulation game vs control	0.83	−1.83, 3.48
Behavior simulation game vs control	2.48	−0.20, 5.15
Attitude/behavior simulation game vs control	2.22	−0.44, 4.88

NOTES: Attitude simulation game = simulation game including role playing/group dynamics.
Behavior simulation game = simulation game including modeling/training.
Attitude/behavior simulation game = simulation game including role playing/group dynamics, and modeling/training.
Control = no intervention.

superior to the means of the groups exposed to behavior triggering elements (Table 1). In addition, there is a significant difference in performance between the group having used the game containing attitude triggering elements and the control group (Table 2).

Transfer of Learning

The three intervention groups obtained a higher mean than the control group on the transfer of learning test (Table 1). However, as shown in Table 2, the confidence intervals of the differences are wide. The group that participated in the simulation game containing behavior triggering elements (modeling and training) scored slightly higher than groups exposed to simulation games containing attitude triggering elements.

Discussion

All three intervention groups did better than the control group on all three tests (attitude, behavior, and transfer of learning). More specifically, the findings on attitude confirm empirical research related to attitude modification (theories of reinforcement,[2–4] group dynamics,[8,9] dissonance,[6,7] and inoculation.)[5] However, contrary to indications from empirical research on attitude modification, the behavior triggering elements (modeling/training) were found to be as effective as the attitude triggering elements since exposure to any of the three simulation games resulted in good traffic safety attitude.

Use of behavior triggering elements in a simulation game appeared to motivate cautious behavior (Table 1). This finding supports those of Bandura[10,11] as well as those of Parcel and Baranowski[12] which indicate that the presence of behavior triggering elements is sufficient to modify behavior. Attitude triggering elements (role-playing/group dynamics) were more effective in motivating cautious behavior than behavior triggering elements themselves (Table 2). The main difference noted is between the subjects using games containing essentially attitude triggering elements and the control group. These results coincide with the studies of Insko,[21] Zimbardo and Ebbesen[22] which indicate that the presence of attitude triggering elements is necessary and sufficient to modify behavior.

Finally, review of the literature on simulation games[13–16] suggests that they are inherently conducive to transfer of learning because they give learners a chance to test their knowledge and skill in an environment similar to reality. Although our findings support those of empirical research (all experimental groups scoring higher than the control group on transfer of learning), our results apply only to five-year-old children and to a simulation game on traffic safety situations. Therefore, they should not be generalized to other populations or to triggering elements other than those contained in the simulation games studied.

In summary, although, our findings do not allow us to conclude that the simulation games will transfer these good traffic safety habits to a real pedestrian situation, we can state that simulation games using triggering elements may serve as an effective educational strategy in modifying attitudes and behavior concerning traffic safety.

ACKNOWLEDGMENTS

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Childhood Injury Deaths: National Analysis and Geographic Variations

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Abstract: Twenty-three causes of injury mortality in children ages 0–14 in the United States were analyzed by age, race, sex, and state of residence for the years 1980–85. Motor vehicles caused 37 per cent of all injury-related deaths and were the leading cause of injury mortality in every group except children younger than one year, for whom homicide was the leading cause. Male death rates were at least four times female rates for suicide, unintentional firearm injury, and injuries related to farm machinery or motorcycles. The drowning rate among Whites was almost twice that of Blacks for ages 1–4, but in the 10–14 year age group the drowning rate for Blacks was over three times that of Whites. In general, the highest injury death rates were in the mountain states and the south. Between 1980 and 1985, the suicide rate in the 10–14 year age group more than doubled. (*Am J Public Health* 1989; 79:310–315.)

Introduction

Injury is the leading cause of death among children above nine months of age in the United States. Injuries cause approximately 44 per cent of all deaths in children ages 1–4, 51 per cent at ages 5–9, and 58 per cent at ages 10–14.[1] The importance of injury deaths in children and young adults is reflected in the fact that injury surpasses all major disease groups as the leading cause of premature years of life lost in this country.[2–5]

We examined 23 categories of fatal childhood injury in the United States, analyzing age, sex, race, and state of residence for the six years 1980–85.

Methods

Mortality tapes for the years 1980 through 1985 were obtained from the National Center for Health Statistics (NCHS). Deaths of children ages 0–14 years, where the underlying event was injury (i.e., coded E800–E999), were analyzed. The 1982 Census Bureau population estimates, multiplied by six, were used as the denominators for calculating the average annual death rates.

Injury deaths were categorized into 23 causes of death based on ICD E-codes. The *expected* number of deaths for each state, based on the death rate for the United States as a whole, was calculated by multiplying the six-year population estimate for the state by the total US death rate for each injury cause. For example, to calculate the expected number of motor vehicle occupant deaths in Alabama, the six-year population estimate of 5,538,000 (the 1982 estimate of 923,000 times six) was multiplied by 3.17/100,000/year (the US motor vehicle occupant death rate) to get 176 expected deaths for 1980 through 1985. As injury death is a relatively rare event in this population, a Poisson distribution was assumed, and 99 per cent confidence limits were calculated for the number of *observed* deaths in each state. In Alabama, the observed number of motor vehicle occupant deaths was 268 with a 99 per cent confidence interval ranging from 226 to 310. If the *expected* number of deaths fell outside the confidence interval set around the *observed* number of death, the state's observed number of deaths was determined to be either significantly high or significantly low. For Alabama, the expected number of motor vehicle occupant deaths, 176, fell well below the lower 99 per cent confidence limit of 226 so Alabama had a significantly high number of observed motor vehicle occupant deaths.

When a state's *observed number of deaths* was significantly high or low, the state was so depicted on a map. In addition, any state whose death *rate* ranked in the top 10 state death rates for a given cause but whose expected deaths fell within the 99 per cent confidence interval of the observed deaths (due to small numbers) were depicted in a separate group called "top ten". All other states were depicted as "average". Sometimes all of the states with the top 10 death rates from a given cause also had significantly high observed numbers of deaths in which case no "top ten" group appears on the map.

Results

National Analysis

During the six-year period 1980–85, almost 10,000 US children ages 0–14 years died of injuries *each year* (Table 1). Thirty-seven per cent of these deaths were related to motor vehicles; this figure includes bicyclist and pedestrian deaths involving motor vehicles. The other leading causes of death were drowning (14 per cent of injury deaths), house fires (12 per cent), and homicide (10 per cent).

When motor vehicle deaths were separated into the subcategories shown in Table 1, the five leading causes of injury death were motor vehicle occupant, drowning, pedestrian, house fire, and homicide. These five causes, in varying order, were consistently the top five in each age, sex, and racial group, with two exceptions: for children less than one year old, suffocation and food aspiration replaced drowning and pedestrian deaths; and for children ages 10–14, bicycles replaced house fires among the top five causes of injury death.

Age: US children ages 0–14 years experienced an average annual death rate of 19.3/100,000 for all injury causes combined during the years 1980–85. The injury death rate was 33.9/100,000/year for children less than one year old, 25.4 for ages 1–4, 14.4 for ages 5–9, and 16.2 for ages 10–14. The relative importance of specific causes of death varied with age (Figure 1). Motor vehicle-related injuries were the largest category for all ages except children less than one year old, for whom homicide accounted for even more deaths than motor vehicles.

The overall death rate from motor vehicles varied less among age groups than most other causes of injury, ranging from 5.5/100,000/year below one year of age to 7.8 for ages 1–

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TABLE 1—Injury Deaths by Cause, Ages 0–14, US, 1980–85

Injury Cause	Number of Deaths 1980–85	Average Annual Death Rate per 100,000 Population
Motor vehicle-related, total	22,174	7.2
Motor vehicle occupant	9,796	3.2
Pedestrian, traffic	7,489	2.4
Bicycle	2,194	0.7
Pedestrian, non-traffic	1,276	0.4
Motorcycle	690	0.2
Other motor vehicle-related	729	0.2
Drowning	8,568	2.8
House fire	7,021	2.3
Homicide	5,722	1.9
Suffocation	2,279	0.7
Firearm, unintentional	1,700	0.6
Fall	1,332	0.4
Aspiration, food	1,248	0.4
Suicide	1,214	0.4
Aspiration, other	983	0.3
Poison, solid/liquid	539	0.2
Farm machinery	508	0.2
Electric current	432	0.1
Poison, vapor	384	0.1
Airplane crash	348	0.1
Other unintentional	3,473	1.1
Unknown intent	1,237	0.4
Medical/surgical	539	0.2
All injury (E800–E999)	59,711	19.3

4 years. Differences emerged, however, when motor vehicle deaths were separated into the six subcategories.

Occupant deaths accounted for 90 per cent of motor vehicle-related deaths prior to age 1 and were also the largest subcategory for ages 1–4 (43 per cent) and 10–14 (46 per cent), but pedestrian deaths were paramount at ages 5–9 (47 per cent). Pedestrian non-traffic deaths, which involve a vehicle striking a child in a driveway or on other private property, comprised 17 per cent of motor vehicle-related deaths in children ages 1–4 years. Annually, an average of 212 children died this way, 83 per cent of them 1–4 years of age.

Bicyclist deaths accounted for 12 per cent of the motor vehicle-related deaths in the 5–9 year age group, and 16 per cent for ages 10–14. Sixty per cent of all childhood bicyclist deaths occurred among 10–14 year old children.

Approximately 115 motorcyclist deaths occurred each year, primarily in the 10–14 year age group. Sixty-seven per cent of these children were identified as the driver of the motorcycle. Some deaths in which the vehicle was classified as a motorcycle may have involved miscoded off-road vehicles, such as all-terrain vehicles and minibikes, which should have been included in the category "other motor vehicle."

Sex: The male-to-female ratio of death rates was 1.7 to 1 for all injury causes combined. Male death rates were over four times the female rates for suicide, unintentional firearm injury, and injury related to motorcycles or farm machinery. For males ages 10–14, unintentional firearm deaths were surpassed only by deaths related to motor vehicles and drowning.

Race: Among the four major racial groups, injury death rates were lowest for Orientals and Whites and highest for Native Americans (Table 2). Causes of injury with exceptionally high death rates for Native Americans were pedestrian non-traffic deaths (3.9 times the rate for all children), poisoning by solid or liquid (3.5 times), aspiration of non-food materials (2.6 times), and motor vehicle occupants (2.2 times).

For Oriental children, the death rate from falls was more than twice the rate for all US children. In contrast, the death rate among Oriental children for unintentional firearm deaths was one-eleventh the US rate and their motorcycle death rate was one-seventh the US rate.

Differences between Blacks and Whites in overall death rates decreased with increasing age. The rate for Blacks was 2.5 times the rate among Whites for children less than one

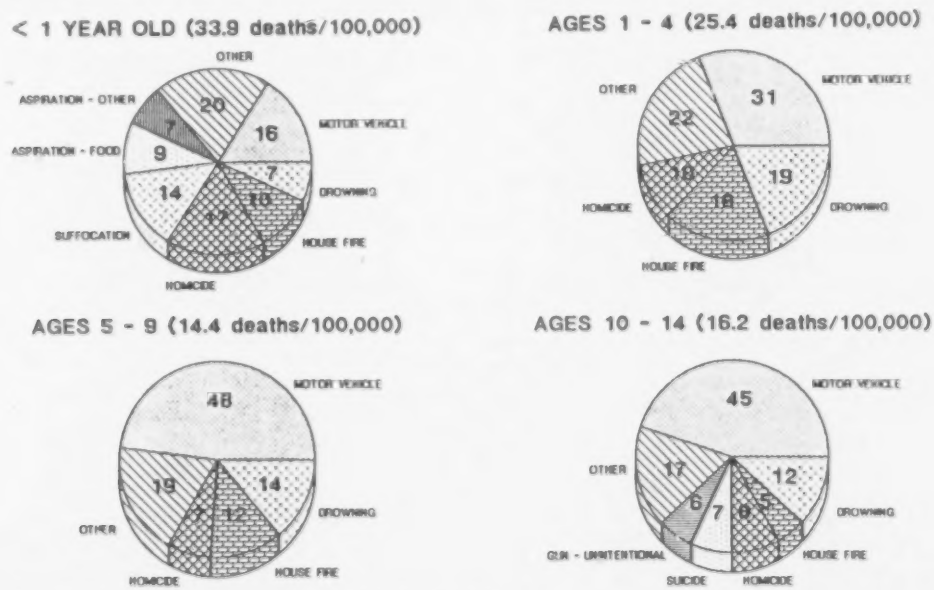

FIGURE 1—Childhood Injury Deaths, Percentage by Cause, US 1980–85

TABLE 2—Average Annual Death Rates* by Leading Causes and Race, Ages 0–14, US, 1980–85

Injury Cause	White	Black	Native American	Oriental	All Races†
All Injuries	17.5 (44,217)	29.3 (13,634)	33.9 (918)	17.5 (917)	19.3 (59,711)
Motor vehicle occupant	3.3 (8,224)	2.6 (1,217)	6.9 (186)	3.1 (162)	3.2 (9,796)
Pedestrian, traffic	2.2 (5,487)	3.8 (1,748)	3.6 (98)	2.9 (153)	2.4 (7,489)
Drowning	2.6 (6,438)	3.9 (1,814)	4.8 (131)	3.4 (180)	2.8 (8,568)
House fire	1.6 (4,080)	6.0 (2,795)	3.3 (90)	1.0 (51)	2.3 (7,021)
Fall	0.4 (933)	0.7 (333)	0.6 (17)	0.9 (49)	0.4 (1,332)
Homicide	1.3 (3,355)	4.7 (2,191)	2.7 (72)	1.9 (101)	1.9 (5,722)

*Rate/100,000/year (total deaths over six-year period in parentheses)
†All Races column includes cases where race was unknown.

year old, 1.8 times for ages 1–4, 1.7 for ages 5–9, and 1.2 for ages 10–14.

House fire and homicide rates for Blacks were more than three times the rates for Whites in all age groups below 10 years. Pedestrian death rates were more than twice as high in Blacks as in Whites for ages 1–4 and 5–9 years. In contrast, the drowning rate among Whites ages 1–4 was almost double that of Blacks. At ages 10–14 years, the most pronounced racial difference was the high death rate from drowning in Blacks (3.4 times the rate for Whites). Black children ages 10–14 continued to experience high homicide and house fire death rates compared to Whites, while White children in this age group had substantially higher death rates than Blacks for motor vehicle occupants, suicide, unintentional firearm deaths, motorcycles, and suffocation.

Time Trends: The injury death rate for the US population ages 0–14 decreased by 16 per cent between 1980 and 1985, from 21.6 to 18.2/100,000/year (Table 3). The greatest decreases occurred in the youngest children: 25 per cent for children less than 1 year old, 21 per cent for ages 1–4, and 19 per cent for ages 5–9, compared to 7 per cent for ages 10–14.

Death rates declined for almost all causes of injury during this time period (Table 3). Among the leading causes of fatal injury, the greatest decreases were seen in rates of drowning and pedestrian deaths. The motor vehicle occupant death rate declined by 14 per cent for all ages combined, with dramatic differences among age groups. Children under 1 year experienced a 32 per cent decrease, ages 1–4 years a 22 per cent decrease, ages 5–9 years an 11 per cent decrease, and ages 10–14 only a 3 per cent decrease. The suicide rate for ages 10–14 more than doubled.

Geographic Variations

Analysis at the state level found more than a three-fold difference among the states with the highest and lowest injury death rates. Injury death rates for all children ages 0–14 ranged from 11.2/100,000/year in Massachusetts to 35.0/100,000/year in Alaska. In general, the highest rates were in the mountain states and in the south, and the lowest rates were in New England, the mid-Atlantic states, and the midwest.

In every state, total motor vehicle deaths surpassed deaths from any other injury cause. Furthermore, in 26 states, the subcategory of motor vehicle occupants outranked deaths from any other category (Appendix). Deaths due to house fires outnumbered those from any other specific injury cause in 11 states, while drowning was the leading injury cause of death in 10 states. Pedestrian deaths ranked first in three states, as did homicide in Washington, DC.

Motor vehicle occupant deaths (Figure 2) were high in the south and southwest. The highest rates were in Idaho (7.4/100,000/year), New Mexico (6.6), and Wyoming (6.5), which

TABLE 3—Death Rates,* Number of Deaths and Per Cent Change in Death Rates, 1980–85, by Cause of Injury, Ages 0–14, US

Injury Cause	1980	1985	Percent Change in Rates
All Injury	21.61 (11,084)	18.19 (9,449)	−16
Suicide (ages 10–14)	0.76 (139)	1.61 (275)	+112
Homicide	1.84 (944)	1.86 (967)	+1
Poison-gas, vapor	0.10 (52)	0.10 (54)	0
Airplane crash	0.12 (60)	0.11 (55)	−8
House fire	2.42 (1,240)	2.15 (1,118)	−11
Firearm-unintentional	0.62 (316)	0.54 (278)	−13
Motor vehicle-occupant	3.60 (1,846)	3.10 (1,610)	−14
Bicycle	0.79 (407)	0.68 (353)	−14
Suffocation	0.77 (393)	0.64 (335)	−17
Pedestrian-traffic	2.84 (1,456)	2.24 (1,161)	−21
Fall	0.48 (245)	0.38 (198)	−21
Aspiration-other	0.37 (191)	0.28 (144)	−24
Drowning	3.60 (1,671)	2.43 (1,263)	−25
Pedestrian-non-traffic	0.48 (248)	0.36 (187)	−25
Farm machinery	0.19 (98)	0.14 (72)	−26
Poison-solid, liquid	0.21 (106)	0.15 (80)	−29
Electric current	0.17 (85)	0.12 (64)	−29
Motorcycle	0.27 (140)	0.19 (98)	−30
Aspiration-food	0.53 (274)	0.35 (180)	−34

*Rate/100,000/year
Number of Deaths in parentheses.

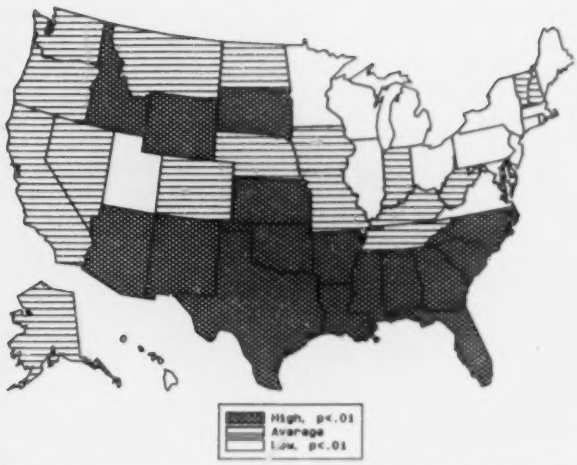

FIGURE 2—Motor Vehicle Occupant Deaths by State, Ages 0–14 Years, US 1980–85

contrasted sharply with the relatively low rate in nearby Utah (2.4).

House fire deaths (Figure 3) showed a distinct geographic pattern, with high rates in the east, especially the southeast, and low rates in the west.

Drowning rates (Figure 4) ranged from 0.8/100,000/year in Rhode Island to 6.8 in Alaska. California, Oregon, and Utah are among the states with significantly high numbers of drowning deaths, although they do not have significantly high numbers of deaths for most other causes.

Death rates from unintentional firearm injury ranged from zero in Rhode Island to 2.1/100,000/year in Alaska. The rates were highest in the south and in the northern mountain states (Figure 5).

Farm machinery death rates were highest in northern farm states: Idaho, South Dakota, Wisconsin, Vermont, Iowa, and Minnesota had rates ranging from 0.6 to 0.9/100,000/year. Deaths from farm machinery had the highest rates among 1–4 year olds and occurred predominantly (80 per cent) among White males. In Wisconsin children ages 1–4, farm machinery deaths were surpassed only by deaths due to motor vehicles, house fires, and drowning (data not shown).

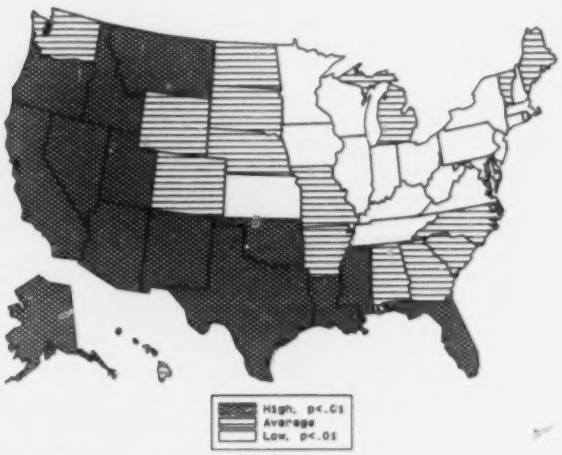

FIGURE 4—Drowning Deaths by State, Ages 0–14 Years, US 1980–85

Homicide rates ranged from 0.7/100,000/year in New Hampshire to 3.0 in Delaware and 4.6 in Washington, DC. High homicide rates were seen throughout the southwest as well as in states with large urban populations (Figure 6).

Suicide rates (Figure 7) are mapped for the 10–14 year age group because only 2 per cent of the childhood suicides occurred in younger children. Suicide rates were generally highest in the northern and mountain states; they ranged from a low of 0.3/100,000/year in North Dakota to 2.1 or higher in Wyoming (3.3), New Mexico (2.7), Idaho (2.4), and Colorado (2.1).

Discussion

Childhood injury death rates have been decreasing for decades[3,6] and it is encouraging that both the frequency and the rate of death from all injuries combined have continued

FIGURE 3—House Fire Deaths by State, Ages 0–14 Years, US 1980–85

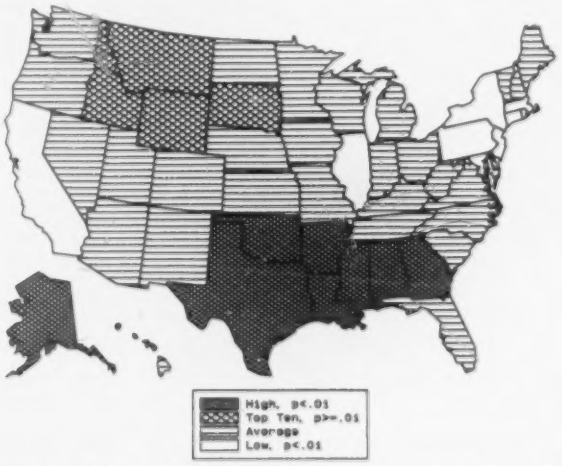

FIGURE 5—Unintentional Firearm Deaths by State, Ages 0–14 Years, US 1980–85

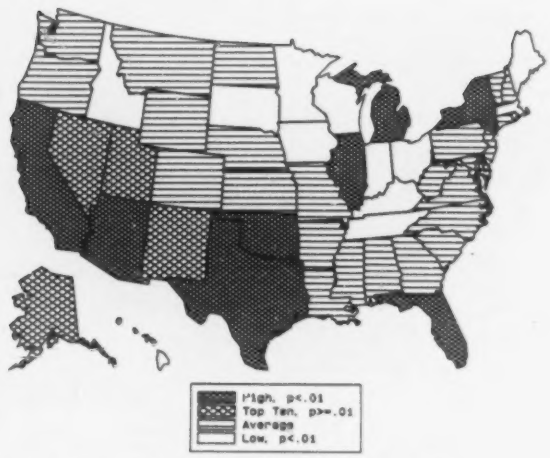

FIGURE 6—Homicides by State, Ages 0–14 Years, US 1980–85

to decrease in recent years. Major causes of unintentional injury death continue to be motor vehicle occupants, pedestrians, drowning, and house fires. The importance of homicide during the first year of life—often representing lethal cases of child abuse—and the increasing rates of juvenile suicide are matters of special concern because they reflect personal, family, and social disruption.

Identifying states with the highest rates of a specific injury cause of death may suggest a likely causal factor. For example, the high house fire death rates in the south and east suggest that the older and/or wooden houses, more commonly found in these regions[3,7] may be a contributing factor. In addition, the National Fire Data Center suggests that low income and education levels and large numbers of families living in poverty, particularly prevalent in the southeast, may contribute to high fire death rates.[8] Similarly, an especially low death rate may be revealing. For example, the low rate of motor vehicle occupant deaths in Utah suggests a beneficial effect on children of social policies (reflecting, in this instance, religious tenets) that reduce alcohol use by adults. The motor vehicle occupant death rate per billion vehicle miles is also low in Utah[3] as well as the per capita consumption of alcohol, which is the lowest in the country.[9]

It is also important to look at the rankings of causes *within* a state. In Maine and Virginia, the house fire death rate is not substantially higher than in the US as a whole—yet house fires are a major problem in these states, with death rates even higher than the rates for motor vehicle occupants. This type of information can be helpful to a state interested in targeting its own most serious problems.*

Limitations of this study include possible coding variations among states. It is unlikely that different coding practices affected the major causes of injury death, since motor vehicle-related injury, drowning, house fire, and homicide are clearly defined and recognizable throughout the country. However, coding of some less prominent injury causes may vary. For example, mechanical suffocation and aspiration, to which about one-third of all injury deaths in children under one year of age are attributed, may be confused with each other and may include miscoded cases of Sudden Infant Death Syndrome (SIDS). More accurate coding of SIDS in recent years may, in part, account for the decrease in deaths coded to aspiration in children less than one year old.

The use of data from a six-year period made it possible to examine causes of death and population subgroups that are rarely considered. For example, pedestrian deaths in non-traffic locations such as driveways and parking lots comprise only 2 per cent of all injury deaths, but 17 per cent of all motor vehicle-related deaths in 1–4 year olds. An examination of pedestrian deaths for this age group in Washington State indicates that fatalities occurring in non-traffic situations may often be misclassified as traffic-related.[10] From 1979 to 1983, one-third of the pedestrian traffic deaths to children under five years of age in Washington State were re-classified as non-traffic deaths when additional information from different data sources was considered. Thus, non-traffic pedestrian deaths may be an even more important cause of death for young children than national data indicate.

The differences observed among racial groups indicate the need for prevention efforts focused on the particular injury problems of certain subpopulations. Falls cause 5 per cent of injury deaths in Oriental children, compared to only 2 per cent in all other children. Falls from or out of buildings are overrepresented among fall deaths in Oriental children when compared to fall deaths in children of other races, suggesting that an urban living environment may contribute to high fall death rates in Oriental children.[11] Data from Korea, where injury mortality patterns are generally quite similar to the US but where the death rate from falls in children ages 1–4 years is over seven times the US rate, suggest a possible cultural difference in the perception of falls as a threat to children's lives.[12]

The high injury death rates in Native American children, as well as the predominance of males among all causes of injury death, were also found in a recent study of childhood injury mortality in Canada.[13] The finding that race differentials in total injury mortality rates decrease with increasing age is consistent with previous research, as is the observation that large race differences still exist for specific injury causes within age groups.[14] For example, the drowning rate for ages 1–4 is highest in White children, probably because of their greater exposure to swimming pools, but for ages 10–14 is far higher for Blacks. Socioeconomic differences have repeat-

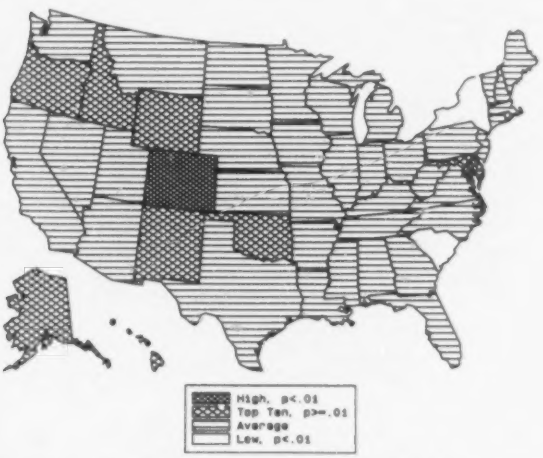

FIGURE 7—Suicides by State, Ages 0–14 Years, US 1980–85

*Detailed data for each state are available from the authors.

edly been shown to be related to differences in injury occurrence as well as injury prevention understanding and behaviors.[15–17] The racial differences revealed by this study most likely reflect differences in income, education, and other aspects of socioeconomic status among racial groups in the US.

This study indicates that in certain geographic areas or for specific population subgroups it is possible to identify injury problems that are often masked in larger, less specific categories. Such identification increases the likelihood that specific injury problems will be targeted and given higher priority in prevention programs, eventually leading to a reduction in injury mortality.

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APPENDIX

Leading Cause of Injury Death* by State, Ages 0–14, US 1980–85

Motor Vehicle Occupant	House Fire	Drowning	Pedestrian	Homicide
Alabama	Delaware	Alaska	Massachusetts	District of Columbia
Arkansas	Illinois	Arizona	New York	
Colorado	Maine	California	Rhode Island	
Connecticut	Maryland	Florida		
Georgia	Michigan	Hawaii		
Idaho	New Jersey	Montana		
Indiana	Pennsylvania	Nevada		
Iowa	South Carolina	Oregon		
Kansas	Vermont	Utah		
Kentucky	Virginia	Washington		
Louisiana	West Virginia			
Minnesota				
Mississippi				
Missouri				
Nebraska				
New Hampshire				
New Mexico				
North Carolina				
North Dakota				
Ohio				
Oklahoma				
South Dakota				
Tennessee				
Texas				
Wisconsin				
Wyoming				

*When all categories of motor vehicle-related deaths are combined, Total Motor Vehicle is the leading cause in *every* state.

Incidence, Causes and Severity of Injuries in Aquitaine, France: A Community-Based Study of Hospital Admissions and Deaths

L. Tiret, PhD, B. Garros, MS, P. Maurette, MD, V. Nicaud, MS, M. Thicoipe, MD, F. Hatton, MD, and P. Erny, MD

Abstract: This paper reports the results of a study of injuries conducted during a one-year period within a defined geographic population of 2.7 million persons (Aquitaine, France). Cases were defined as unintentional or intentional injuries, either resulting in immediate death before reaching hospital or requiring hospital admission. During the one-year period, 1,181 deaths were registered and 8,190 hospital admissions occurred during the sample periods. The three leading causes of injury were falls (40 per cent), traffic accidents (27 per cent), and poisonings (15 per cent). The overall incidence of injuries was 136 per 10,000 person years. Incidence by sex and age was assessed for the main external causes. The ratios of nonfatal to fatal cases were calculated by sex and age and by external cause. The origin of the injury was suicide in 14 per cent of cases and assault or homicide in 3 per cent. The severity of injuries, assessed using an automatic computation of the Injury Severity Score (ISS), ranged from 1 to 66 with a mean of 6.9. Substantial variations of ISS were observed according to external cause. At the 8th day following admission, 31 per cent of hospital-treated patients were still hospitalized and 0.8 per cent had died in hospital. The outcome correlated well with the ISS. (*Am J Public Health* 1989; 79:316–321.)

Introduction

Injuries constitute a major health problem in all developed countries and in an increasing number of developing countries. They have been examined by mortality data[1,2] or by morbidity data focusing either on the pediatric population[3–7] or on specific injury types, such as head traumas,[8,9] burns,[6,10] or motor vehicle injuries.[7,11] Very few studies describe the nature of the problem as a whole within an entire population. In some countries, routine statistics[12,13] or permanent surveys, such as the National Health Interview Survey in United States,[14] provide annual estimates of the incidence of injuries but without great detail. More information was gathered by the Northeastern Ohio Trauma Study,[15,16] carried out in 1977 in the United States. However, that study did not include fatal cases and did not provide an evaluation of the severity of injuries.

This paper reports the first results of an epidemiological study of injuries conducted during 1986 among a defined geographic population in France. Its aims were to provide estimates of the incidence of fatal and nonfatal cases of injury, to describe their external causes, and to assess the severity of injuries using the Injury Severity Score (ISS) derived from the Abbreviated Injury Scale (AIS),[17] the most widely used score for rating severity of multiple injuries.[18–21]

Methods

The study was carried out prospectively over a one-year period in the Aquitaine, one of the 22 administrative regions of France. This region contains both urban and rural areas; its main city is Bordeaux. It is bordered by the western Atlantic coastline and the southern mountain range of Pyrénées. Extrapolating from the 1982 census, the 1986 population was estimated to be 2.7 million persons (4.9 per cent of the whole population of France), of which 16 per cent were over 65 years of age. The study was based on an administrative region since in the French health care system, each region constitutes an independent entity and people normally receive medical care in their region of residence. The presence of a referral teaching hospital in the region reduces the possibility of transfer for specialized treatment outside the region.

Case Definition

Cases were defined as those injuries sustained in Aquitaine between December 1985 and December 1986, either resulting in immediate death before reaching hospital, or requiring hospital admission in a private or public institution of the region. Injuries to residents and non-residents of the region were both recorded but the latter were excluded from the data reported here. The field of study included both unintentional and intentional injury cases (the latter including homicides, assaults, suicides, and self-destructive actions). Surgical and medical misadventures (corresponding to codes E870–E879 and E930–E949 of the International Classification of Diseases—9th revision[22]) and late effects of accidental injury (E929) were excluded.

Data Collection

All public and private hospitals regularly involved in the treatment of injuries, because of their geographical position or the presence of a trauma specialist, were included in the study. Among other hospitals, one out of two was randomly selected. In each institution, the survey lasted 13 weeks spread over the whole year:

- the teaching hospital was investigated every fourth week;
- non-teaching and private hospitals of the first category were divided into four groups, each being investigated during three periods (4 weeks, 4 weeks, and 5 weeks). The survey was initiated successively in each group with an interval of four weeks, the whole year therefore being covered by the study;
- private hospitals of the second (non-trauma) category were divided into two groups, each being successively investigated during a 13-week period.

This scheme provided a sample representative by time of day, day of the week, and season.

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Injury cases resulting in death before reaching hospital were recorded from death certificates. All cases registered in the region during the one-year period were included, except for those having been registered in a hospital or a clinic.

Data Collected

For each injury case requiring hospital admission, the medical staff of the institution completed a questionnaire within the first 24 hours following admission. Data collected concerned the demographic characteristics, time and place of injury, origin, cause and clinical nature of injury. Causes of injury and poisoning were coded into 14 specific categories. Cases with two or more possible causes were coded according to the primary harmful event. Outcome within eight days of injury was added later. The clinical nature of traumatic injuries was coded according to a classification specially devised for this study by a multidisciplinary workgroup, including trauma specialists and epidemiologists. The purpose of this classification was to provide an automatic computation of the ISS. The descriptive lesions, taken directly from the AIS-80 manual,[17] were condensed into 200 codes being organized within the six body regions defined for the ISS. A defined injury code corresponded either to one specific lesion or to a group of several lesions of a same anatomic site and having an identical AIS severity level. The person in charge of completing the questionnaire simply had to circle the codes relative to all injuries sustained, then an AIS grade was automatically assigned to each injury code by a computer algorithm and the ISS was computed.*

A test to control the adequacy of ratings by this method was undertaken on 164 patients treated in an emergency department. One physician computed the ISS using the manual. The same lesions were then coded in the questionnaire and the ISS was derived by the algorithm. In 14 per cent of cases, there were slight discrepancies between the two ratings, due to the adaptation, which over-simplified some categories of lesions in the AIS, such as spinal lesions.

The completeness of the notification of cases was checked by two supervisors for each period of survey, from the list of admissions to the institutions.

Analysis

Patients of unknown age and sex were excluded from the analysis (1 per cent of total). In the event of transfer, only the first hospital admission for the episode was considered. Data concerning hospital admissions were extrapolated from the sample to provide annual estimates. To improve the accuracy of estimated injury rates, an attempt was made to adjust for the missing cases diagnosed outside the region. This adjustment was made on the basis of the 1984 mortality statistics.[23] Among deaths from injury having occurred in residents of Aquitaine, the proportions of cases registered outside the region were calculated by sex and age group, and the estimated injury rates were increased by the same proportions, on the assumption that admissions and deaths follow the same pattern. This proportion varied from 3.9 per cent to 10.7 per cent (6.3 per cent for the entire population).

Tables concerning the ISS were derived only from the admission sample, since this information was not available for death cases. ISS was analyzed both as a continuous variable and in classes, those being the categories widely used.[19,24]

*Details about the methodology are available from the author on request.

Results

All 21 public hospitals and 38 of the 43 private hospitals participated in the study. The participating institutions accounted for 97 per cent of the total traumatological activity in the region. The coefficient of extrapolation from the sample of admissions was 4.09.

Incidence

During the one-year period, 1,181 deaths were registered from the death certificates and 8,190 hospital admissions were observed during the sampling periods among residents of Aquitaine. After extrapolation of admissions on an annual basis and adjustment for the potential cases diagnosed outside the region, the overall incidence was 135.9 per 10,000 person years.

As expected, injury rates varied considerably by sex and age (Table 1). Men between 15 and 24 years of age and women over 75 years of age experienced the highest overall injury rates; except for ages 65 and over, male rates exceeded female rates, the difference being especially marked in the 15–34 year age group. However, mortality rates for males markedly exceeded those for females at all ages except those under age 10. In both sexes, mortality rates were highest over 75 years of age.

There was one immediate death for every 30 hospital admissions. The rate ratio of nonfatal to fatal cases was very high during childhood, then decreased until age 65 for females and age 75 for males. Except under age 10, the ratio was lower in males than in females.

External Cause of Injury

For all ages combined, falls on same level were the most common cause of fatal and nonfatal injuries (Table 2). When adding falls from high level, falls represented nearly 40 per cent of all reported cases of injury, yielding an incidence of 52 per 10,000 person years. Traffic accidents were the second commonest cause, followed by poisonings. These three leading causes accounted for 80 per cent of all admissions and deaths. When considering only fatal cases, the rank ordering of causes differed markedly. Traffic accidents ranked first, followed by accidents caused by firearms, hanging and strangulation, and drowning. The cause of injury was unspecified in 8 per cent of death certificates.

The nonfatal to fatal case rate ratio varied considerably with external cause of injury. For example, there was one immediate death in 150 admissions from falls on same level, against one in 20 from traffic accidents. In cases of drowning and hanging, the number of immediate deaths exceeded that of admissions. The ratio was also very low for injuries caused by firearms or electricity and very high for injuries produced by blunt objects and cutting or piercing instruments.

Origin of Injury

Among fatal and nonfatal cases of injury, 13.9 per cent were suicides or attempted suicides and 2.7 per cent were assaults or homicides (Table 3). Suicides and attempted suicides accounted for more than 80 per cent of cases of hanging/strangulation and poisonings, more than half of injuries due to firearms, and one-third of drownings. It should be noted that for this last cause, the origin was unknown in nearly 10 per cent of cases. More than 25 per cent of injuries from firearms were of accidental origin, hunting being a common sport in Aquitaine. Two suicides by electricity were reported.

TABLE 1—One-year Injury Rates per 10,000, by Sex and Age, Aquitaine, France, 1986

Age (years)	Sex	Hospital Admissions		Deaths		Overall Annual Rate*	Non-fatal/Fatal Rate Ratio
		No. in Sample	Extrapolated Annual Rate*	Annual No.	Annual Rate*		
<5	M	239	131.2	8	1.1	132.3	119.3
	F	165	91.3	7	1.0	92.3	91.3
5–9	M	225	120.2	5	0.7	120.9	171.7
	F	131	71.0	5	0.7	71.7	101.4
10–14	M	300	142.4	12	1.4	143.8	101.7
	F	159	77.3	2	0.2	77.5	386.5
15–24	M	1,244	268.4	122	6.5	274.9	41.3
	F	549	116.8	42	2.2	119.0	53.1
25–34	M	833	185.4	141	7.7	193.1	24.1
	F	386	82.7	33	1.8	84.5	45.9
35–44	M	578	147.1	104	6.5	153.6	22.6
	F	354	85.3	47	2.8	88.1	30.5
45–54	M	382	121.8	106	8.3	130.1	14.7
	F	258	75.3	39	2.8	78.1	26.9
55–64	M	389	109.5	130	9.0	118.5	12.2
	F	290	74.9	52	3.3	78.2	22.7
65–74	M	208	90.8	89	9.5	100.3	9.6
	F	286	103.3	40	3.5	106.8	29.5
≥75	M	298	172.4	104	14.7	187.1	11.7
	F	916	283.4	93	7.0	290.4	40.5
All ages	M	4,696	157.4	821	6.7	164.1	23.5
	F	3,494	106.7	360	2.7	109.4	39.5
	Both	8,190	131.3	1,181	4.6	135.9	28.4

*Rates were adjusted for the potential cases diagnosed outside the region.

Cause of Injury by Sex and Age

For both sexes, falls and traffic accidents ranked among the three leading causes at any age. Women over 75 years of age experienced the highest rate of falls on same level. Men were more exposed than women to falls from a high level, except during childhood where rates were similar in both sexes. Traffic accidents showed a marked peak in the 15–24 year age group, with a male/female ratio of incidence in this group of 2.5. Two distinct poisoning risk groups emerged: the first under 5 years of age, corresponding to accidental cases, and the second between 15 and 44 years of age, more marked in females and corresponding primarily to intentional poisonings. Injuries produced by blunt objects and cutting/piercing instruments showed an excess of male cases mainly concentrated during the working ages. Burns were more frequent in males at any age, and in both sexes, the highest risk was experienced under 5 years of age (Table 4).

Injury Severity Score (ISS)

Among patients having sustained traumatic injuries, the ISS ranged from 1 to 66 with a mean of 6.9. Eleven per cent of patients had an ISS higher than 10. Accidents caused by firearms yielded the highest values of both mean and proportion of ISS above 20. Traffic accidents and burns were also associated with a mean ISS value above average. Falls on same level often led to a score of 9, reflecting the fractures of limbs. In contrast, falls from high level were associated with higher proportions of scores in extreme classes, reflecting

TABLE 2—One-year Injury Rates per 100,000, by External Cause of Injury, Aquitaine, France, 1986

Cause of Injury	Hospital Admissions		Deaths		Overall Annual Rate*	Non-fatal/Fatal Rate Ratio
	No. in Sample	Extrapolated Annual Rate*	Annual No.	Annual Rate*		
Falls on same level	2366	379.4	63	2.5	381.9	151.8
Traffic accidents	2168	347.6	391	15.3	362.9	22.7
Poisonings	1202	192.7	69	2.7	195.4	71.4
Falls from high level	851	136.4	34	1.3	137.7	104.9
Struck by object	515	82.6	10	0.4	83.0	206.5
Cutting/piercing	509	81.6	6	0.2	81.8	408.0
Burns	121	19.4	11	0.4	19.8	48.5
Firearms	53	8.5	166	6.5	15.0	1.3
Insect/animal bite/sting	75	12.0	4	0.2	12.2	60.0
Drowning	25	4.0	137	5.4	9.4	0.7
Hanging/strangulation	12	1.9	159	6.2	8.1	0.3
Foreign body	41	6.6	5	0.2	6.8	33.0
Electricity	10	1.6	8	0.3	1.9	5.3
Miscellaneous	130	20.8	19	0.7	21.5	29.7
Unspecified	112	18.0	99	3.9	21.9	4.6

*Rates were adjusted for the potential cases diagnosed outside the region.

TABLE 3—Origin of Injury according to External Cause

External Cause of Injury	Annual Extrapolated Number*	Origin (%)			
		Suicide	Homicide Assault	Other	Unspecified
Traffic accidents	9260	0.1	—	99.2	0.3
Falls on same level	9740	—	1.3	98.5	0.2
Falls from high level	3510	2.4	0.2	97.0	0.4
Drowning	240	30.2	—	60.2	9.6
Foreign body	170	—	—	97.6	2.4
Struck by object	2120	0.6	25.0	73.4	1.0
Cutting/piercing	2090	5.9	5.2	87.8	1.2
Firearms	380	53.7	17.3	26.7	2.4
Burns	500	5.8	0.8	93.2	0.2
Poisonings	4990	80.7	0.2	18.0	1.0
Hanging/strangulation	210	89.7	8.3	1.0	1.0
Electricity	50	20.8	—	79.2	—
Insect/animal bite/sting	310	—	5.3	94.7	—
Miscellaneous	550	0.5	3.2	96.3	—
Unspecified	560	4.9	2.2	50.8	42.2
All causes	34,680	13.9	2.7	82.2	1.2

*Admissions × 4.09 + deaths.

accidents of different severity. Being struck by an object or cutting/piercing from instruments produced the least severe injuries (Table 5).

Outcome of Patients Admitted to Hospital

On the eighth day following injury, 31 per cent of patients admitted to hospital were still hospitalized and less than 1 per cent had died in hospital. When controlled for age by using the Standard Morbidity Ratio (SMR), burns, firearm accidents, and traffic accidents had proportions of patients still hospitalized above average. Hospital mortality was higher than average for traffic accidents, firearms and hanging, and lower for poisonings (Table 6).

The outcome within eight days correlated well with the ISS, mortality after admission and length of hospitalization both increasing with the score (Table 7).

Discussion

In contrast to studies based on medical records,[3,4,15] our data were specifically collected for this study, using an ad hoc questionnaire. This system of data collection was time-consuming for the medical staff of institutions in charge of completing questionnaires but, in return, data were more comprehensive and relevant than those derived from routine medical records.

This study does contain a number of possible biases. First, the field of study was limited to severe cases, i.e., cases either resulting in immediate death or requiring hospital admission. Hospital admission may be influenced by several factors independent of the actual severity of injury such as age, social conditions, access to medical service, hospital admission policy. Nevertheless, this criterion was no more subjective than the fact of attending at an accident and emergency department.

Concerning the sample of admissions, five of 43 private hospitals did not participate in the study. The extrapolation from the sample was adjusted for these missing data, taking surgical activity as criterion of reference although this did not exactly reflect the traumatological activity.

In participating institutions, completeness of notifications was checked by two supervisors using the list of admissions, but this list was not available in all private hospitals and some underreporting of cases in these institutions cannot be excluded. However, private institutions accounted for only 25 per cent of all admitted cases.

The number of deaths before admission may also be underestimated, since in case of a "suspect" cause of death, some certificates may have been sent directly to the legal authorities. These cases are later re-included in official mortality statistics, but they often have an unspecified cause of death, since legal authorities do not always send back the relevant information. In the 15–34 year age group, where it can be assumed that unspecified causes correspond mainly to violent deaths, we can predict from official mortality statistics that we may have missed about 9 per cent of cases. This

TABLE 4—One-year Injury Rates* per 10,000 by External Cause of Injury, Sex and Age, Aquitaine, France, 1986

Age (years)	Sex	Falls on Same Level	Traffic Accidents	Poisonings	Falls from High Level	Struck by Object	Cutting Piercing	Burns
0–4	M	28.9	13.8	28.5	29.5	4.6	8.3	8.3
	F	20.7	11.1	19.2	25.8	3.9	2.8	3.3
5–9	M	41.3	25.8	3.8	23.1	10.2	5.4	1.6
	F	22.0	10.5	2.7	24.7	7.0	2.7	1.2
10–14	M	51.0	42.8	2.9	17.6	13.4	6.8	1.4
	F	26.2	21.6	7.8	16.5	2.4	2.4	0.5
15–24	M	35.4	127.8	25.3	16.7	28.8	22.6	4.1
	F	10.4	49.6	43.2	5.4	1.7	5.0	1.3
25–34	M	25.5	58.8	25.0	20.8	20.0	21.8	2.9
	F	5.9	23.1	43.4	3.5	1.5	4.6	1.5
35–44	M	23.5	37.8	20.4	19.4	17.4	15.3	2.8
	F	9.8	23.8	37.9	4.5	2.0	3.2	1.5
45–54	M	17.8	34.1	9.8	24.5	13.1	9.4	2.2
	F	14.4	24.5	23.4	5.6	2.6	2.9	1.6
55–64	M	23.2	33.2	6.2	15.6	10.5	13.5	0.9
	F	36.1	15.5	11.0	4.2	1.5	1.0	0.5
64–74	M	32.9	31.5	2.9	13.4	3.6	4.9	1.7
	F	62.8	14.0	6.1	9.0	3.2	1.4	0.7
≥75	M	106.3	31.8	10.1	16.5	4.2	4.0	1.7
	F	235.6	15.5	6.2	9.8	2.2	2.5	1.4
All ages	M	33.7	50.7	15.1	19.1	14.7	13.8	2.7
	F	42.2	23.0	23.6	8.9	2.4	3.0	1.3

*Rates were adjusted for the potential cases diagnosed outside the region.

TABLE 5—Classes of Injury Severity Score (ISS), Mean and Range according to External Cause of Injury

External Cause of Injury	Classes of ISS (%) 1–4	5–9	10–19	≥20	Mean (S.D.)	Minimum Maximum
Traffic accidents	34.1	43.7	15.7	6.5	8.2 (0.16)	1–66
Falls on same level	38.5	57.4	3.5	0.6	6.7 (0.07)	1–36
Falls from high level	44.8	42.6	10.6	2.0	7.0 (0.16)	1–34
Struck by object	62.5	32.0	4.3	1.2	5.0 (0.20)	1–41
Cutting/piercing	73.6	22.8	3.0	0.6	4.8 (0.14)	1–25
Firearms	49.1	24.5	3.8	22.6	10.2 (1.66)	1–50
Burns	51.2	33.8	10.0	5.0	8.0 (0.62)	4–29
All traumatic causes	43.7	44.9	8.4	3.0	6.9 (0.07)	1–66

underestimation presumably affected suicides and homicides.

With respect to the estimates of injury rates, we had to assume that fatal and nonfatal cases followed the same pattern for potential cases diagnosed outside the region, an assumption which may not be accurate.

Lastly, hospital mortality possibly is underestimated, since in cases of a secondary transfer to another institution, the occurrence of death in the secondary institution was not known. However, this underestimation is doubtless slight (only 4 per cent of cases were secondarily transferred). Hospital mortality assessment was limited to eight days, because it did not seem feasible to maintain a longer follow-up in all institutions. This limitation does not affect equally all types of injury, delayed deaths being more frequent for some causes and ages.

The comparison of our results with those of other studies is difficult since many differences exist in definition and case ascertainment, time and place of study, and analysis. Nevertheless, many findings are similar to those reported by previous studies,[3,4,12–16] concerning the main causes of injury and the patterns of incidence by age and sex.

The major contributions of this study lie in two areas. First is the assessment of the ratio of nonfatal to fatal cases.

TABLE 6—Outcome at the Eighth Day following Injury of Patients Admitted to Hospital, according to External Cause of Injury

External Cause of Injury	Number of Cases	Still Hospitalized %	Still Hospitalized SMR	Died in Hospital %	Died in Hospital (Nb)	Died in Hospital SMR
Traffic accidents	2168	31.3	1.16*	1.4	(31)	1.94*
Falls on same level	2366	44.4	1.01	0.8	(20)	0.71
Falls from high level	851	30.8	1.02	0.8	(7)	1.02
Drowning	25	24.0	0.91	—		—
Foreign body	41	10.3	0.40°	—		—
Struck by object	515	16.9	0.62*	0.2	(1)	0.27
Cutting/piercing	509	14.3	0.52*	—		—
Firearms	53	49.1	1.73*	7.5	(4)	9.22+
Burns	121	61.0	2.28*	0.8	(1)	1.23
Poisonings	1202	17.4	0.70*	0.1	(1)	0.16°
Hanging/strangulation	12	41.7	1.57	8.3	(1)	11.56+
Electricity	10	20.0	0.72	—		—
Insect/animal bite/sting	75	11.1	0.38*	1.3	(1)	1.88
Miscellaneous	130	27.9	0.96	—		—
All causes	8190	31.3	1.0	0.8	(67)	1.0

SMR = Standardized morbidity (mortality) ratio. SMR = [Observed number of cases still hospitalized (died) for a given cause/Expected number when applying the overall rate of patients still hospitalized (died)]. *$p < 0.001$; °$p < 0.05$; +expected number smaller than 5.

TABLE 7—Outcome at the Eighth Day following Injury of Patients Admitted to Hospital, according to Classes of ISS

Classes of ISS	% Still Hospitalized	% Died in Hospital
1–4	17.7	0.1
5–9	42.2	0.5
10–19	58.5	1.1
≥20	69.9	18.5

The ratio is very high during childhood, because hospital admissions are frequent and mortality very low. From 15 to 75 years of age, the ratio regularly decreases since rate of fatal cases increases, while that of nonfatal cases decreases. In the elderly, the ratio rises again, especially in women, despite a continuing increase of immediate mortality. This is explained by the high admission rates at this age and by the fact that deaths are more often delayed, as mentioned above. The low ratios of nonfatal to fatal cases for certain external causes, such as accident by firearms, drowning, and hanging/strangulation, emphasize the fact that studying severe cases only from hospital-treated cases does not provide an accurate view of the magnitude of the problem.

The second contribution of this study is an in-depth description of the severity of injuries using the ISS. We have not found such an evaluation of the ISS within an entire injured population in the literature. The ISS was initially developed for traumas resulting from traffic accidents, but according to its authors, its use for rating other accidents is possible on condition that similar injuries are involved.[18] Moreover, the 1980-AIS version incorporated many more injuries to internal organs, making the AIS more suitable for rating severity of non-vehicular penetrating injuries.[21] Its application may be questionable for rare causes, such as firearm accidents.[25]

We developed a methodology for automatically computing the ISS, as proposed by other authors.[26] The purpose was to eliminate the inter- and intra-rater variability among coders, a problem which has been often pointed out.[21,25–27] This problem was important in our study, which involved many institutions. Automatic computation was also time-saving and less costly, two necessary conditions to make the study feasible, even at the expense of some loss of precision. MacKenzie,** reporting the results of an experiment of conversion of the International Classification of Diseases to AIS, had already stimulated interest in such a tool, even not perfect.

The ISS could be assessed only on patients admitted to hospital, since in cases of death a postmortem examination would have been necessary to identify all lesions. This minimizes the proportion of upper scores for causes having a high immediate mortality, such as traffic accidents or firearms injuries. Despite this bias, these causes were nevertheless associated with the highest severity scores among hospital-treated patients.

This study confirmed that the ISS correlated well with mortality and length of hospitalization, as previous studies have reported.[18,19] However, it is not a good predictor of

**McKenzie EJ, Steinwachs DM, Shankar B: Classifying trauma severity based on hospital discharge diagnoses: validation of an ICD-9CM to AIS-85 conversion table (not published).

post-trauma disability.[24] Further exploration of this last aspect is needed as the improvements in survival from injuries may lead to an increase in later disability.

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APHA Issues New AIDS Report:
Contact Tracing and Partner Notification

The American Public Health Association's Special Initiative on Acquired Immune Deficiency Syndrome (AIDS) recently released its second report, entitled *Contact Tracing and Partner Notification*. This new report, which was prepared by the Association's AIDS Working Group, presents the historical and scientific basis for contact tracing and partner notification in the control of sexually transmitted diseases. The report also discusses those factors which should be considered before deciding if contact tracing is a desirable public health strategy in preventing the transmission of the human immunodeficiency virus (HIV) which causes AIDS.

Single copies of *Contact Tracing and Partner Notification* are available for $3 prepaid from APHA Publication Sales, 1015 15th Street, NW, Washington, DC 20005. Bulk rates are also available on request. Telephone (Publication Sales): (202) 789-5666.

The first report prepared by the APHA AIDS Working Group, released last year, was entitled *Casual Contact and the Risk of HIV Infection*. That 12-page report presents the epidemiologic and scientific evidence regarding the major routes of HIV, the rare routes of HIV infection, and those routes that do not transmit HIV. Single copies of *Casual Contact and the Risk of HIV Infection* are available for $3 prepaid from APHA Publication Sales (address and phone number above). Bulk rates also available.

Sport Fish Consumption Advisories and Projected Cancer Risks in the Great Lakes Basin

JEFFERY A. FORAN, PHD, MICHAEL COX, MPH, AND DAVID CROXTON, MPH

Abstract: Because Great Lakes sport fish are contaminated with several toxicants, the Great Lakes states individually issue advisories, principally based on Food and Drug Administration (FDA) action levels, that suggest limiting or eliminating consumption of contaminated fish. We describe the procedures the states use to determine when to issue consumption advisories and we evaluate the associated cancer risks using EPA-IARC-OSTP risk assessment procedures. Projected cancer risks are high for consumers of small quantities of sport fish contaminated with DDT or dieldrin at their respective action levels. Projected risks at concentrations that are common but below the action levels are also substantial. We propose that sport fish with tissue concentrations of DDT or dieldrin one-fifth and one-third of the action levels should be covered by consumption advisories to warn consumers of the potential adverse health impacts. (*Am J Public Health* 1989; 79:322–325.)

Introduction

Contamination of the sport fishery in the Great Lakes basin (Lakes Michigan, Huron, Superior, Erie and Ontario and their connecting waters) has been a highly sensitive environmental and human health problem. Since the early 1970s, states in the Great Lakes basin (Illinois, Indiana, Michigan, Minnesota, New York, Ohio, Pennsylvania, Wisconsin) individually have issued annual consumption advisories that suggest reduction or elimination of the consumption of contaminated sport fish. Following the Great Lakes Toxic Substances Control Agreement, states were partially successful in implementing coordinated advisories in 1986. These advisories are based on Food and Drug Administration (FDA) action levels for commercial fish and vary substantially among the states. Whether the consumption advisories protect the population is not known.

In this report, we describe the different consumption advisories, which are dependent upon tissue concentrations of contaminants in local fish populations and we use quantitative cancer risk assessment to evaluate their efficacy.

Methods

We calculated risk projections for two contaminants of Great Lakes sport fish, DDT (1,1,1,-trichloro-2,2-bis[p-chlorophenyl]ethane) and dieldrin (1,2,3,4,10,10-hexachloro-6,7-epoxy-1,4,4a,5,6,7,8,8a-octahydro-endo-exo-1,4,:5,8-dimethanonaphthalene), partly because of recent regulatory action on these materials.[1,2]

Projected risks were calculated following the guidelines of the US Environmental Protection Agency (EPA),[3,4] Office of Science and Technology Policy (OSTP),[5] and the International Agency for Research on Cancer (IARC).[6] We used data from the studies of mammalian toxicity for DDT and dieldrin conducted by Cabral, *et al*,[7] Tomatis, *et al*,[8] and Walker, *et al*.[9] We used the multistage model[10] (Global 82) for extrapolation from high to low doses and rodent to human scaling factors. For each combination of fish consumption and contamination level, we computed four cancer risk estimates by using either rat or mouse experimental data and either surface area or body weight scaling to estimate 95 per cent upper bound risk projections.

Concentrations of DDT (expressed as total DDT) and dieldrin in the edible tissue of Great Lakes sport fish (coho, chinook and king salmon, and lake trout) varied widely throughout the Great Lakes during 1986 (0.01 mg/kg (ppm) to 1.5 mg/kg for total DDT and 0.01 mg/kg to 0.2 mg/kg for dieldrin, Michigan Department of Natural Resources and Wisconsin Department of Natural Resources, unpublished data). Both DDT and its metabolite DDE are found in Great Lakes sport fish, with long range atmospheric transport acting as a continuing source of DDT to the Great Lakes. We chose intermediate tissue concentrations of DDT (total) and dieldrin (1.0 mg/kg and 0.1 mg/kg) to represent those found in many Great Lakes sport fish in 1986 and to be used for comparisons with FDA action levels. The FDA action level for DDT in fish is 5.0 mg/kg (ppm) and 0.3 mg/kg for dieldrin.

Three levels of consumption used in the risk assessment range from the US national average (6.5 g/day or less than one portion per month weighing one half pound)[11] to 96 g/day (approximately three, half pound portions per week) over a lifetime. Surveys suggest that the average consumption rate of sport fish by anglers and their families in the Great Lakes basin ranges from 10 to 50 g/day[12–14] but that some individuals may consume sport fish at least three times per week (96 g/day).[15]

Results

States surrounding Lake Michigan issue consumption advice for sport fish based on the number of fish sampled with edible tissue concentrations that exceeded the FDA action level for individual contaminants (Table 1). States surrounding the other Great Lakes generally issue advice based on the mean tissue concentration for individual contaminants. Consumption advice is modified by the states annually and is issued on a location, species and size-class specific basis.

Cancer risk projections (Table 2) range as high as 1.05×10^{-2} (one excess occurrence of cancer during the lifetimes of 100 exposed individuals living to age 70) for dieldrin and as high as 2.8×10^{-3} for DDT when consumption is 96 g/day (three portions per week) and fish are contaminated with DDT and dieldrin at their respective action levels (5.0 mg/kg, 0.3 mg/kg).

When 10 per cent to 50 per cent of sampled Lake Michigan sport fish are contaminated with DDT or dieldrin at or above action levels, consumers are advised to reduce consumption of

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TABLE 1—Process for Developing Sport Fish Consumption Advisories for the Great Lakes (excludes Ontario)

Lake	Advice		
	Unrestricted Consumption (Category 3)	Restrict Consumption* (Category 2)	Do Not Eat (Category 1)
Michigan	0–10% of fish sampled exceed Action Level	11–50% of fish sampled exceed Action Level	>50% of fish sampled exceed Action Level
Huron	Mean Tissue Concentration < Action Level	Mean Tissue Concentration > Action Level	Mean Tissue Concentration > 2X Action Level
Erie	Mean Tissue Concentration < Action Level	Advice Not Given	Mean Tissue Concentration > Action Level
Superior	variable**	variable**	Mean Tissue Concentration > Action Level
Ontario	None	All fish[a] Mean Tissue Concentration 1.5–3 X AL[b]	Mean Tissue Concentration > 3X AL[b]

Note: All advice is based on the concentration of contaminants in edible tissue of individual sport fish species or size classes of individual species.

*States advise that women of child bearing age, pregnant women, and children do not eat any fish in this category.

**Advice for these categories is based on FDA Action Levels but the procedure for applying action levels is highly variable within and between states.

[a]New York issues general advice that all fish should not be consumed more than once per week.

[b]This advice is based on an additivity formula where concentrations of more than one contaminant in fish tissue are combined. Where this combination of contaminants is 1.5–3X the combined action level, New York advises that these fish should not be eaten more than once per month.

these species and size classes of fish to not more than one portion per week. When more than 50 per cent of sampled fish contain contaminants above the action level, no consumption is recommended. Cancer risk projections associated with consumption of one portion per week (32 g/day) of fish with tissue concentrations at action levels range as high as 9.4×10^{-4} for DDT and as high as 3.8×10^{-3} for dieldrin (Table 2).

Although concentrations of DDT and dieldrin in the edible tissue of Great Lakes sport fish were below action levels in 1986, cancer risk projections associated with tissue concentrations of 1.0 ppm DDT or 0.1 ppm dieldrin range as high as 5.7×10^{-4} for DDT and 3.7×10^{-3} for dieldrin. At a fish consumption rate of 32 g/day, risk projections associated with these tissue concentrations range as high as 2.0×10^{-4} for DDT and 1.5×10^{-3} for dieldrin.

All of the projections cited above were derived using a

TABLE 2—95 Per Cent Upper Bound Cancer Risk Projections in Humans Exposed to Fish Contaminated with DDT or Dieldrin

Fish Consumption (g/d)*	SCL[a]	ANM[b]	Tissue Concentration (PPM)			
			DDT		Dieldrin	
			1.0 PPM	5.0 PPM†	0.1 PPM	0.3 PPM†
6.5	BW	M	3.6×10^{-6}	1.5×10^{-5}	4.51×10^{-5}	9.02×10^{-5}
		R	1.7×10^{-6}	7.0×10^{-6}		
	SA	M	4.8×10^{-5}	2.0×10^{-4}	5.86×10^{-4}	1.04×10^{-3}
		R	9.5×10^{-8}	3.9×10^{-5}		
32.0	BW	M	1.5×10^{-5}	7.1×10^{-5}	1.13×10^{-4}	2.93×10^{-4}
		R	7.2×10^{-6}	3.3×10^{-5}		
	SA	M	2.0×10^{-4}	9.4×10^{-4}	1.49×10^{-3}	3.76×10^{-3}
		R	4.0×10^{-5}	1.9×10^{-4}		
96.0	BW	M	4.3×10^{-5}	2.1×10^{-4}	2.93×10^{-4}	7.89×10^{-4}
		R	2.0×10^{-5}	9.9×10^{-5}		
	SA	M	5.7×10^{-4}	2.8×10^{-3}	3.65×10^{-3}	1.05×10^{-2}
		R	1.1×10^{-4}	5.5×10^{-4}		

*32 g/day = 1/2 pound/week

†FDA action level.

[a]SCL - BW = body weight scaling factor, SA = surface area scaling factor

[b] ANM - M = mouse, R = rat

rodent to human scaling factor based on surface area. A body weight scaling factor reduces risk projections, generally by an order of magnitude for any consumption level and tissue concentration. Use of a body weight scaling factor results in risk projections greater than 1×10^{-5} for all consumption rates of dieldrin and greater than 1×10^{-6} for DDT (Table 2). Use of a body weight scaling factor results in risk projections as high as 7.9×10^{-4} for dieldrin and 2.1×10^{-4} for DDT when fish contaminated at the action level are consumed three times per week. Risks associated with consumption of fish contaminated at current levels range as high as 2.9×10^{-4} for dieldrin and 4.3×10^{-5} for DDT at a consumption rate of 96 g/day. Thus, while the use of a body weight scaling factor results in lower cancer risk projections, these projections are still high for all but the lowest consumption rates and tissue concentrations.

Finally, the use of toxicological data derived from a rat study for DDT results in differences in risk projections of less than an order of magnitude compared with risk projections derived from toxicological data from the mouse study. This is true for comparisons within any consumption level and tissue concentration.

Discussion

Our projections suggest that current consumption advisories for Great Lakes sport fish are inadequate and that consumers of those fish may face substantial excess cancer risks when tissue concentrations of DDT or dieldrin in fish are at or near the FDA action level. They may even face excess cancer risks at one-third to one-fifth of the action level if they consume fish at least weekly. By implication, advisories to restrict or avoid consumption may be necessary when tissue concentrations are one-third to one-fifth of the FDA action levels.

Several caveats apply to our modeling of risk. First, the problems inherent in extrapolation from rodent experiments to humans, the use of the multistage model, the use of benign or malignant tumors, species, sex and organ differences in tumor development rates, and the fact that DDT and dieldrin may act as promoters in the formation of cancer[16–21] may all result in inaccurate predictions of human cancer rates. The US EPA states that actual cancer rates may be much lower than the 95 per cent upper bound estimates and may actually be as low as zero. We have followed the EPA-IARC-OSTP guidelines for handling these specific points. Models are currently under development to incorporate methods of action as well as metabolic, physiologic and biochemical considerations,[22,23] but they have not been adopted or endorsed by the US EPA or FDA.

A second problem is the inconsistent data from more than 30 animal studies on the carcinogenicity of DDT, DDE and dieldrin. Even among the few studies that meet EPA-IARC-OSTP guidelines for acceptable cancer studies, not all report positive results. On the other hand, it is reassuring that the use of rats or mice and the use of DDT or DDE changes risk projections by less than one order of magnitude.

Third, the model used here may underestimate cancer risk by ignoring concurrent exposure to other contaminants.[24] Great Lakes sport fish carry several chlorinated organic contaminants in their tissues.

Fourth, the risk assessment model used here does not incorporate consideration of any benefits that may be associated with fish consumption such as reduced blood cholesterol levels and reduced risk of heart attack. Nor does it consider that proper cleaning and cooking procedures may reduce the level of some, but not all, contaminants (for example, cooking does not reduce mercury levels in fish tissue).

Finally, the risk assessment model used here does not address other toxicological endpoints. Exposure to chlorinated organic toxicants have been shown to affect visual recognition memory (PCBs),[25] motor immaturity and other neonatal endpoints (PCBs),[26] duration of breast feeding (DDE),[27] and the occurrence of chronic lymphatic leukemia (DDT).[28]

We propose that a standard, risk assessment based approach be adopted for the development of fish consumption advisories. This approach should be used to project cancer risks and combined with risk management decisions such as acceptable levels of risk and the level of acceptable economic impact, if any. This approach is appropriate until epidemiological data are available to validate current risk assessment models or until improved modeling techniques are adopted that incorporate consideration of all human health impacts as well as pharmacokinetic information, risk associated with exposure to multiple contaminants, and any risk/benefit trade-offs that are associated with consumption of sport fish.

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On Job/On Campus Program at U-MI SPH for Fall 1989

Public Health Policy and Administration

The University of Michigan School of Public Health will offer a new on job/on campus master's level program in "Public Health Policy and Administration" beginning in October 1989. The non-residential program features monthly sessions over an extended weekend on campus for a two-year period—modeled after the School's successful 16-year program in "Health Services Management."

The new program will focus on the development and enhancement of skills and knowledge needed for effective public health leadership in state and local public health and community health agencies. Designed for upper and mid-level managers in public health agencies as well as other community health leaders who need to expand their skills and knowledge in policy analysis, formation, and implementation, the program will emphasize integration of work experience with classroom sessions using lectures, cases, seminars, projects, and computer conferencing. Faculty will be drawn from the U-MI School of Public Health and other schools and colleges at the University, in addition to experts from both public and private institutions and programs.

For further information on this program, contact George Pickett, MD, or John Romani, PhD, Department of Public Health Policy and Administration, U-MI SPH, Ann Arbor, MI 48109-2029. Tel: (313) 764-2132.

Occupational Medicine and Industrial Hygiene

A similar program in "Occupational Medicine and Industrial Hygiene" at the U-MI School of Public Health will be offered for the fifth time in the Fall of 1989. Applications are now being accepted in the work-study program in occupational health leading to a master's degree in two years. Students attend campus once per month for three-day sessions that extend from Thursday morning through Saturday afternoon.

The Occupational Medicine and Industrial Hygiene Program includes course work in biostatistics, epidemiology, toxicology, occupational diseases, ergonomics, and industrial hygiene. For further information, contact David Garrabrant, MD, or Thomas Robins, MD, OJ/OC Occupational Health Program, School of Public Health, U-MI, Ann Arbor, MI 48109-2029. Tel: (313) 764-2594.

Public Health Briefs

Driver Age and Crash Involvement

ALLAN F. WILLIAMS, PHD, AND OLIVER CARSTEN, PHD

Abstract: The youngest and oldest drivers have the highest crash risk, but the problem lies predominantly in the youngest age groups because elderly drivers have low exposure. The elderly driver problem will increase gradually as their share of the population increases but will remain relatively small. The bulk of the problem will continue to reside among drivers younger than age 65, particularly the youngest drivers. (*Am J Public Health* 1989; 79:326–327.)

Introduction

As the average age of the population of the United States increases, attention is being given to the economic, social, and health problems that will be created or aggravated by there being more older people. The problem of motor vehicle injuries has received increased attention in this regard because the oldest drivers, along with the youngest, are known to have high crash risks when they drive.[1,2] In this paper, current relationships between age and driver crash involvement are described, and an assessment is made of the effect an aging population will have on overall driver crash rates and the contribution of various age groups to the problem during the next 40 years.

Method

The data used to compute crash involvement rates and to estimate future trends were as follows:

- Estimates of miles driven were obtained from the most recent (1983) National Personal Transportation Study (NPTS), based on a national survey of 6,438 households containing 17,382 members selected so as to represent the total civilian noninstitutional population of the United States.[3] As part of the survey all members of sampled households were asked about their travel on a randomly selected day. The data used here have been modified from the public use data supplied by the Federal Highway Administration. In the survey, many household trips were reported by someone other than the driver, and, if for some reason the driver was not subsequently interviewed, no driver trip was recorded. For such cases (about 9 per cent of all passenger car trips), no information on driver age and sex was available from the public use tapes. In the present study, the household trip data were reexamined and all driver trips are included.
- Numbers of drivers of passenger vehicles (cars, light trucks, vans) in fatal crashes were obtained from the Fatal Accident Reporting System (FARS), a computerized data base containing information on virtually all motor vehicle fatalities in the United States.[4]
- Estimated numbers of drivers of passenger vehicles in all police-reported crashes were obtained from the National Accident Sampling System (NASS), a probability sample of police-reported crashes in the United States.[5] To reduce variability from small sample sizes, three years of NASS data, 1982–84, were combined. The original national sampling weights were used, but the final estimate was divided by three to give an annual count.
- Population data (historical, present, and projected population estimates) were obtained from the US Bureau of the Census.[6–8]

Results

The youngest and oldest drivers have the highest crash rates per miles driven, both for fatal and for all crashes (Figure 1). In terms of total numbers of drivers in fatal crashes, however, the problem lies predominantly in the youngest age groups (Figure 2). Half of all drivers in fatal crashes in 1983 were 16–29 years old.

The relatively small number of elderly drivers involved in fatal crashes despite their high crash risk is attributable to three main factors: there are fewer persons per single year of age among the elderly than among younger age groups, the rate of licensure is lower among the elderly, and licensed elderly drivers drive fewer miles on average than do younger drivers (Table 1).

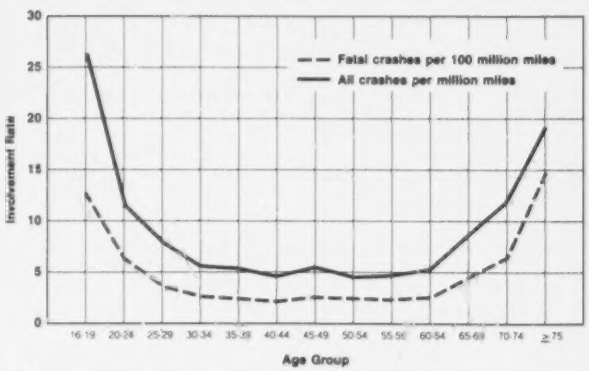

FIGURE 1—Mileage-Based Crash Rates by Driver Age, United States, 1983

Address reprint requests to Allan F. Williams, PhD, Insurance Institute for Highway Safety, Watergate 600, Washington, DC 20037. Dr. Carsten is with the Institute for Transport Studies, University of Leeds. This paper, submitted to the Journal December 24, 1987, was revised and accepted for publication October 24, 1988.

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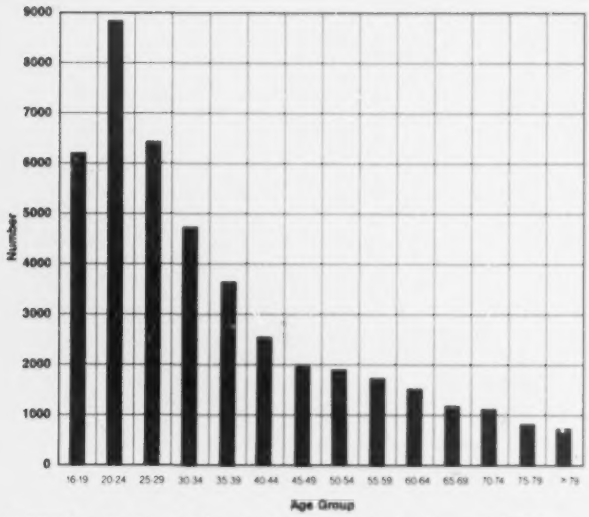

FIGURE 2—Number of Drivers in Fatal Crashes by Age Group, United States, 1983

In 1987, the 65-and-older group comprised 16 per cent of the population and 10 per cent of the drivers in fatal crashes (16.3 per 100,000 population). In contrast, 16 to 19 year-olds contributed 8 per cent of the population and 14 per cent of drivers in fatal crashes (47.1 per 100,000); 20 to 29 year-olds were 22 per cent of the population and 34 per cent of the drivers in fatal crashes (40.5 per 100,000).

The 65-and-older population is growing at a faster rate than other age groups; the result is that their population share has been increasing and will continue to do so, but the increase will be gradual until after the year 2010. In 2010, the 65-and-older group will represent 17 per cent of the population over age 15; by 2030, they will form 26 per cent of the population. However, because the age distribution is shifting such that an underrepresented group (the elderly) is becoming a larger proportion of the population and an overrepresented group (the young) is becoming a smaller proportion, if current (1987) per capita rates prevail the overall rate of drivers in fatal crashes would be lower in 2010 (24.5 per 100,000) and in 2030 (23.8) than it was in 1987 (26.5). These projected numbers indicate that in 2010 and 2030 there will be many more elderly drivers in fatal crashes, but in 2030 there would be more than twice as many 16 to 29 year-olds as 65-and-older drivers involved, and about the same number of teenage drivers as drivers over age 70.

Discussion

There has been increasing discussion of loss-reduction measures directed specifically to this age group, such as night driving curfews, special licensing actions, and driver training programs tailored to older drivers.[2,9] It is important, however, to keep the problem of elderly drivers in perspective.

TABLE 1—Per Cent Licensed and Average Annual Passenger Vehicle Miles Driven per License Holder, by Age Group, 1983

Age Group (years)	Per Cent Licensed	Average Miles Driven
16–19	61	5,199
20–29	83	8,529
30–39	92	9,748
40–49	89	9,141
50–59	88	7,679
60–69	79	5,743
70–74	41	4,463
>74	36	2,667

Older drivers have an elevated crash risk when they drive but their exposure is low; their current contribution to the problem is small and, although increasing, will remain relatively small well into the next century. The bulk of the crash problem will continue to reside with drivers younger than age 65, and particularly with the youngest drivers. Identification of effective and feasible loss-reduction measures for high-risk subgroups such as the elderly is obviously important, but these efforts should not detract from the broader goal of development and application of countermeasures aimed at drivers of all ages.

It is possible that the per capita driver fatal crash involvement rate of the elderly will increase more (or decrease less) than that of other age groups in the next decades because, among other reasons, future generations of the elderly may be relatively older or they may drive more miles. If this occurs, the contribution of the elderly will be increased, although young drivers will almost certainly continue to be the predominant group in terms of numbers involved.

ACKNOWLEDGMENT

This work was supported by the Insurance Institute for Highway Safety.

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Unintentional Deaths from Carbon Monoxide in Motor Vehicle Exhaust: West Virginia

ROY C. BARON, MD, MPH, RONALD C. BACKER, PHD, AND IRVIN M. SOPHER, DDS, MD

Abstract: We investigated the circumstances of unintended carbon monoxide deaths from motor vehicle exhaust. Of 64 episodes involving 82 deaths investigated by the West Virginia Office of the Chief Medical Examiner, 1978–84, 50 occurred outdoors in older vehicles with defective exhaust systems and 14 occurred in enclosed or semi-enclosed home garages. Blood alcohol was detected in 50 (68 per cent) of 74 victims tested; 34 had blood alcohol concentrations ≥0.10g/dl. We suggest increasing public awareness of the hazards of motor vehicle exhaust and enforcing vehicle inspection regulations. (*Am J Public Health* 1989; 79:328–330.)

Introduction

Carbon monoxide (CO) from motor vehicle exhaust is the leading cause of unintentional fatal poisoning in the United States.[1] In 1980, it accounted for 611 (14 per cent) of the nation's 4,331 unintended poisoning fatalities,[2] 2.7 deaths per 1,000,000 population. High rates are more common in northern states, in rural regions, and in areas with low per capita income.[1]

In West Virginia, which ranks second among the 50 states in rural inhabitants (63.8 per cent)[3] and 41st in per capita income ($6,141),[4] motor vehicle exhaust caused 76 (33 per cent) of the 231 unintended fatal poisonings among state residents in the years 1978–82,[5] almost three times the expected number based on the national rate in 1980 (95 per cent CI for observed/expected ratio, 2.3, 3.6).

We reviewed all unintentional deaths from CO related to motor vehicles for the years 1978–84 using data based principally from the Office of the Chief Medical Examiner (OCME).

Methods

The OCME in West Virginia has jurisdiction over all homicide, suicide, accidental, and unattended natural deaths that occur in the state and maintains a centralized file of all case records. When indicated, toxicologic analyses for blood alcohol concentration (BAC) and per cent CO saturation are performed by standard methods.[6,7]

We reviewed the OCME files for the seven-year period, 1978 through 1984, to identify all unintended deaths attributed to CO from motor vehicle exhaust. Data recorded included each victim's age, sex, BAC, and per cent CO saturation, along with the date, location, and other circumstances associated with the episode. Specific vehicle defects that might have contributed to the fatal episode were recorded from information either found in the OCME record or acquired through special request from local law enforcement agencies. The vehicles' last previous dates of inspection were not available.

A review of vital statistics for in-state deaths during 1978 (ICD-8 code E873) and 1979 through 1984 (ICD-9 code E868.2) was conducted to identify additional deaths from motor vehicle exhaust, certified as unintentional, but which had not been reported to the medical examiner. These were included only in the tabulations of age, sex, and month of occurrence since blood analyses were not performed and other details were not available.

Results

For the years 1978–84, we identified 68 separate episodes of fatal unintentional exposure to CO from motor vehicles in West Virginia, resulting in 87 deaths. Four episodes (five deaths) had not been reported to the OCME. Sixty-eight (78 per cent) decedents were male. Only one victim was a child, and rates were highest among persons 15–34 years of age, who accounted for 64 per cent of the deaths (Table 1). Episodes and deaths were most common during winter and fall and least common in late spring and summer (Figure 1). With the exception of one pick-up truck and one motorized van, all vehicles were passenger cars.

TABLE 1—Deaths Due to Unintentional Carbon Monoxide Poisoning from Automotive Exhaust, by Victim's Age, West Virginia, 1978–84

Age of Victim (years)	No. Deaths (%)	Rate*
<15	1 (1)	0.3
15–24	34 (39)	14.0
25–34	22 (25)	10.3
35–44	8 (9)	5.5
45–54	8 (9)	5.8
55–64	8 (9)	5.8
65+	6 (7)	3.6
All	87 (100)	6.4

*Deaths per 1,000,000 population per year; computations based on United States Bureau of Census enumeration for 1980.

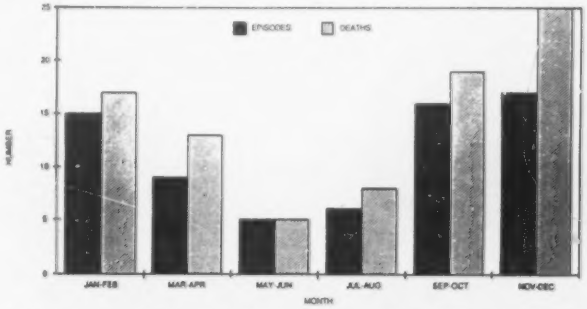

FIGURE 1—Fatal Unintentional Carbon Monoxide Poisoning from Automotive Exhaust, Episodes and Deaths, by Month of Occurrence, West Virginia, 1978–1984

From the Office of Epidemiology and Health Promotion (Dr. Baron) and the Office of the Chief Medical Examiner, (Drs. Backer and Sopher) of the West Virginia Department of Health, and the Division of Field Services, Epidemiology Program Office, Centers for Disease Control (Dr. Baron).

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TABLE 2—Medical Examiner-Investigated Cases of Unintentional Carbon Monoxide Poisoning from Automotive Exhaust: Deaths by Location and Circumstance, West Virginia, 1978–84

Circumstance	Location					
	Outdoors		Garage		Total	
	Episodes	Deaths	Episodes	Deaths	Episodes	Deaths
Moving vehicle	4	4	0	0	4	4
Couples parked	9	17	1	2	10	19
Car repair	0	0	5	5	5	5
Stuck in snowy ditch	3	3	0	0	3	3
Other stationary vehicle	34	43	8	8	42	51*
Total	50	67	14	15	64	82

*Includes 42 persons whose prolonged car occupancy was attributed to alcohol consumption, 3 to smoking marijuana, and 2 each to conversation, use of air conditioning, or unknown.

Of the 64 episodes investigated by the OCME, 50 occurred outdoors and 14 occurred in home garages. Forty-one (76 per cent) of 54 episodes for which the information was available were estimated to have occurred between midnight and 6 am. Although not recorded routinely, car heaters were frequently reported to have been on in episodes during cooler months (September through April). Of 11 episodes that occurred during May through August, the car air conditioner had been on in at least four, and no mention of its use was recorded in the other seven.

All except four vehicles were stationary when the victims were overcome (Table 2). By available history, 42 of the 51 decedents in "other stationary" vehicles had been drinking alcohol at the time of death or were sleeping or resting in their cars after drinking earlier. Their cars were found parked in remote roadside areas (19 deaths), in empty parking lots (six deaths), outside their residences (nine deaths), or in their garages (eight deaths).

The median age of the 50 vehicles involved in outdoor episodes was 10 years (range five to 18), substantially greater than the six year median for all passenger vehicles registered in West Virginia as of July 1982.[8] Information from postmortem inspections, available for 41 of these vehicles, revealed that 40 had defective exhaust systems, with or without other body damage. Specified defects included: rust holes in the tailpipe or muffler, loose connections between the manifold and the exhaust pipe, short or missing tailpipes, and partial obstruction of tailpipes with dirt or snow. Damage noted on the car body included corrosion of the floorboards, side panels, or trunk linings.

Six of the 14 episodes in home garages occurred with the main garage door open and three others occurred with only an access door or a garage window left open. Circumstances in the remaining five garage-related episodes, ruled as accidental by the OCME, suggested that the main door had been left closed inadvertently.

Analyses performed on 74 adult decedents revealed 68 per cent with detectable BAC (Table 3), ranging up to 0.33g/dl. These included 43 (73 per cent) of 59 men and seven (47 per cent) of 15 women. In the majority of persons positive, BACs were ≥0.10 g/dl. The presence of blood alcohol was more common and observed in higher concentrations among persons found in "other stationary vehicles" than among all other decendents combined.

Discussion

Unintentional deaths from CO in motor vehicle exhaust occur year-round in West Virginia, generally in the late evening or early morning hours when people are likely to seek rest, companionship, privacy, or shelter in their automobiles. Consistent with previous reports,[9,10] most of the deaths occurred in older model vehicles that were outdoors and stationary and had badly worn, altered, rusted, or partially obstructed exhaust systems.

A majority of the garage-related episodes occurred despite open garage doors or windows, underscoring the failure of passive ventilation to be a reliable means of removing CO from semi-enclosed spaces. Since the affinity of hemoglobin for CO is approximately 220 times greater than that for oxygen,[11] even low concentrations of CO can be hazardous during prolonged exposure.

Prevention of CO deaths in motor vehicles depends, in part, on increasing public awareness of the hazards of motor vehicle occupancy while the engine is idling, particularly in older model vehicles or when the vehicle is snowbound or in semi-enclosed spaces. Drinking behavior was a predisposing factor. Since much alcohol use among victims of fatal injury is associated with problem drinking,[12] however, more ratio-

TABLE 3—Medical Examiner-Investigated Cases of Unintentional Carbon Monoxide Poisoning from Automotive Exhaust: Blood Alcohol Concentration (BAC) by Circumstance of Death, West Virginia, 1978–84

Circumstance	No. Tested for BAC	No. with Positive BAC		Total (%) Positive
		.01-.09g/dl	>.10g/dl	
Moving vehicle	3	0	1	1 (33)
Couples parked	19	5	2	7 (37)
Car repair	5	0	0	0 (0)
Stuck in snowy ditch	3	1	1	2 (67)
Other stationary vehicle	44	10	30	40 (91)
Total	74	16	34	50 (68)

nal behavior may not be a reasonable expectation. Measures to remove the hazard should be more effective than those requiring modification of behaviors or attitudes that are difficult to change or that require frequent implementation.[13]

All motor vehicles in West Virginia are required by statute to undergo an annual inspection that includes "a complete and thorough check of the exhaust system, including manifolds, gaskets, muffler, exhaust lines, brackets, and clamps." The law lists specific standards and provides for rejection if they are not met. Unfortunately, the law does not encourage compliance by requiring evidence of inspection as a condition for annual registration, and is enforced only by rare roadway spot checks for valid stickers. Furthermore, there is currently no control over the quality of the inspections.

Because we were unable to obtain the interval since last inspection, we cannot specify the number of episodes that might be averted by better compliance. Including such information as part of the routine investigation of unintended deaths from CO in motor vehicles would be important to understand the value of more effective enforcement.

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The Inaccessibility of Seat Belts in Taxicabs

RONALD M. DAVIS, MD, MA

Abstract: The accessibility of seat belts was determined in a total sample of 200 taxicabs waiting to pick up passengers at eight city airports. A rear seat belt was "accessible" (able to be fastened within 10 seconds) in 111 (55.5 per cent) taxicabs. The proportion varied by city from 16.0 per cent (New York City) to 96.0 per cent (Minneapolis-St. Paul) and was higher for taxicabs in cities covered by mandatory state seat belt legislation. (*Am J Public Health* 1989; 79:330–331.)

Introduction

In 1985, there were 1.91 billion passengers in 144,235 taxicabs in the United States.[1] The safety of transportation in taxicabs could be compromised for several reasons. Taxicabs are driven an average of 52,000 miles per year,[1] a distance that might increase the risk of mechanical failure. Because of long hours on the road, taxicab drivers may be prone to fatigue, and economic incentives may encourage taxicab drivers to drive at excessive speeds. Seat belts, if present, may not function properly, or may be less accessible to passengers by falling underneath or behind the seat cushions.

There is anecdotal evidence that seat belts are inaccessible in many taxicabs.[2–4] This article describes an investigation of the accessibility of seat belts in taxicabs in eight cities.

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Methods

From May 1986 through September 1986, I assessed the accessibility of seat belts in taxicabs at a major airport in eight cities: Washington, DC (National Airport); Oklahoma City (Will Rogers World Airport); Atlanta (Hartsfield Airport); Chicago (O'Hare Airport); New York City (LaGuardia Airport); Minneapolis-St. Paul (Minneapolis-St. Paul International Airport); Ft. Lauderdale, Florida (Hollywood International Airport); and Las Vegas (McCarren International Airport). At each airport, taxicabs were randomly selected from among those waiting to pick up customers, until information had been collected for 25 taxicabs. Limousines were excluded.

I positioned myself in the rear seat of each taxicab on the far right-hand side, the most likely position to be occupied by a passenger. A seat belt was defined to be accessible if it could be fastened within 10 seconds. This was assumed to be the maximum length of time that a typical passenger would spend attempting to locate and fasten the seat belt. Seat cushions were not moved to facilitate the extraction of seat belts, but I attempted to grasp seat belts by probing between the seat cushions.

Information was collected from the drivers of the taxicabs on the year that the car was manufactured, the taxicab company, and the availability of an infant car seat for use by passengers. The existence of legislation mandating the use of seat belts among adults was also ascertained.

Miettinen test-based confidence limits were calculated for rate ratios.

Results

Information was obtained for 25 taxicabs in each of the eight cities. Seven taxicab drivers refused to participate in the study (three in Atlanta and one each in four other cities).

TABLE 1—Passenger Seat Belt Accessibility in Taxicabs in Eight Cities, United States, 1986

City	State Seat Belt Legislation*	Date of Survey	No. of Taxicabs	No. of Taxicab Companies	Median Year of Taxicab Manufacture	Taxicabs with Accessible Seat Belt**	
						No.	(%)
Washington, DC	A/NE	May 6, 1986	25	21	1978	16	(64.0)
Chicago	A/NE	June 19, 1986	25	5	1984	16	(64.0)
Minneapolis	A/NE	Aug. 29, 1986	25	17	1984	24	(96.0)
Ft. Lauderdale	A/NE	Sept. 21, 1986	25	3	1984	20	(80.0)
Subtotal			100			76	(76.0)
Oklahoma City	B/NE	May 19, 1986	25	6	1978	15	(60.0)
New York City	A/E	July 29, 1986	25	12†	1984	4	(16.0)
Atlanta	C	May 22, 1986	25	10	1978	10	(40.0)
Las Vegas	C	Sept. 27, 1986	25	7	1984	6	(24.0)
Subtotal			75			20	(26.7)
TOTAL			200			111	(55.5)

*Status of mandatory state seat belt legislation for adults at time of survey: A = law adopted and in effect, B = law adopted but not yet in effect, C = no legislation adopted, E = taxicabs exempted from legislation, NE = taxicabs not exempted from legislation. All existing laws in these cities affect front seat passengers only.

**See text for definition.

†Twelve taxicabs owned by different companies, 13 taxicabs privately owned.

Of the 200 taxicabs sampled, a seat belt was accessible in 111 (55.5 per cent) (Table 1). The proportion varied from 16.0 per cent in New York City to 96.0 per cent in Minneapolis-St. Paul. None of the 200 taxicabs had an infant car seat available.

Taxicabs in the four cities covered by mandatory seat belt legislation were more likely to have an accessible seat belt (76.0 per cent) than taxicabs in cities not covered by legislation (26.7 per cent) (rate ratio = 2.8, 95% confidence intervals = 2.1, 3.9). This difference is potentially confounded by the fact that the cities covered by mandatory seat belt legislation were slightly more likely than other cities to have recently manufactured taxicabs. When the cities with older taxicabs are excluded, however, taxicabs in cities covered by seat belt legislation remained more likely to have accessible seat belts (rate ratio = 4.0, 95% CI = 2.7, 6.0), although the number of cities (five) in this comparison is small.

Discussion

Only about half of all taxicabs in eight urban samples had an accessible seat belt. No data were collected on the reasons for the inaccessibility of seat belts (e.g., absent, nonfunctional, or displaced). However, in most cases seat belt straps were partially accessible or visible but the buckles were trapped behind or under the seat cushions. This suggests that the reduced accessibility of seat belts in taxicabs is caused primarily by the buckles falling (or being placed) behind or under the seat cushions. Several taxicab drivers in Las Vegas attributed seat belt inaccessibility to the fact that some drivers pull out the rear seat cushion at the end of their shift to look for money or gambling chips that passengers may have lost; this practice would allow the belts to fall under the cushions.

In 1984 and 1985, 42 fatalities of taxicab occupants were reported to the National Highway Traffic Safety Administration (Grace Hazzard, personal communication, September 25, 1986). There were 251 total occupants of the taxicabs involved in these fatal accidents. Only one (3.6 per cent) of 28 passengers using a restraint was killed, compared with 29 (21.6 per cent) of 134 passengers not using a restraint (relative risk = 6.1, 95% CI = 1.3, 30.2). (There were 12 fatalities among 89 passengers for whom restraint usage was unknown.) This pattern of increased risk with failure to use restraints is similar to that seen in all fatal motor-vehicle accidents.[5]

In 1985, there were 105.5 total nonfatal injuries and 2.7 serious nonfatal injuries for every fatality among occupants of passenger cars.[5,6] Applying these ratios to the 42 fatalities among taxicab occupants in 1984 and 1985 provides a crude estimate of 4,431 total nonfatal injuries and 113 serious nonfatal injuries among taxicab occupants. Many of these deaths and injuries could have been prevented by high occupant seat belt usage, which obviously requires the ready accessibility of seat belts.

State seat belt laws should apply to taxicabs, and should require that properly functioning seat belts be accessible and that infant car seats be available. By 1990 virtually all new cars will be equipped with rear seat lap/shoulder belts.[7] However, this will not necessarily solve the problem of inaccessible seat belts in taxicabs because the recipient portion of the belt could still fall behind or underneath the seat cushion, and because many pre-1990 models will remain on the road for several years.

Meanwhile, taxicab companies and authorities should provide signs that encourage passengers to use seat belts. All taxicabs sampled in Chicago had the following message posted next to their fares: "Please fasten seat belts: They have been installed for your safety and protection."

ACKNOWLEDGMENTS

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Coverage Gaps in Seat Belt Use Laws

JoAnn K. Wells, BA, Allan F. Williams, PhD, and Michele Fields, JD

Abstract: Some seat belt use laws exempt occupants in certain seating positions and in certain types of vehicles, or both. A state-by-state analysis was conducted of the extent to which passenger vehicle occupants killed before the laws went into effect would have been covered by the laws subsequently passed. Only five states had laws that covered all passenger vehicle occupants; the majority excluded 4 to 8 per cent and six states excluded 14 per cent or more. More inclusive laws can reduce injuries and save lives. (*Am J Public Health* 1989; 79:332–333.)

Introduction

During the four-year period 1984-87, laws requiring motor vehicle occupants to use seat belts were passed in 30 states and the District of Columbia. These laws vary from state to state in requirements, enforcement provisions, and penalties. Most states exempt some occupants from coverage even though those occupants have seat belts available for use. Some of these exemptions apply to only a few occupants (e.g., rural letter carriers and delivery vehicle drivers); others apply to far greater numbers (e.g., occupants of pickup trucks). In other countries, such exemptions have been found to limit the effectiveness of seat belt laws because use rates among the exempted populations do not increase or they increase less than in populations the laws cover.[1,2] A study of child restraint laws enacted in all 50 states between 1978 and 1985 found that nationwide about 20 per cent of the children killed in pre-law years who were within the age limits set by the subsequent laws in their state would not have been covered by the laws.[3] The present study examines the gaps in coverage of belt use laws applicable to older populations.

Method

The belt use laws enacted in 31 jurisdictions as of the end of 1987 were analyzed for coverage provisions. Using the Fatal Accident Reporting System[4] (FARS), data on motor vehicle occupant deaths were tabulated for each jurisdiction for 1980–84, the five years preceding the enactment of state belt use laws. FARS is a computerized data base containing information on virtually all fatal motor vehicle crashes in the United States. The data are collected by the state governments under contract to the National Highway Traffic Safety Administration.

Excluded from the study population were children who had already been covered by child restraint laws, occupants of any type or model year of motor vehicle not required to be equipped with seat belts, occupants in seating positions not required to be equipped with belts, and occupants of large trucks (greater than 10,000 pounds). Large trucks were excluded because, prior to implementation of the belt use law examined in the present study, interstate truckers had been required by federal regulations to use seat belts, and many states required belt use in at least some trucks traveling intrastate. The type of vehicle was obtained from FARS using the code for vehicle type and model year was determined by the variable for model year. Only these occupants in a vehicle with a known model year were used in the analysis. The data were analyzed in terms of provisions of each state law to determine how many of the fatally injured occupants would subsequently have been covered by the seat belt use laws enacted in that state. The numbers of occupants killed who would not have been covered by subsequent state laws were calculated for the various states and according to type of exemption.

Results

Six states excluded light trucks and vans as of August 1988. Twenty-six of the 31 jurisdictions exempt all rear seat occupants other than children (who had already been covered by child restraint laws); three states (Minnesota, Missouri, New York) included in their seat belt law some children older than those who had been covered by the prior child restraint law. Only five states—California, Montana, Nevada, Washington, and Wisconsin—cover all passenger vehicle occupants.

Table 1 shows for each jurisdiction the total number of passenger vehicle occupant deaths in 1980-84, the number who would not have been covered by the subsequent state law because of exemptions for occupants in rear seats or for occupants of light trucks and vans, and the total number and per cent exempted.

In the 31 jurisdictions combined, 8,064 (7.6 per cent) of the motor vehicle occupants killed in 1980-84 would not have been covered by the subsequent state laws; 5,579 (5.3 per cent) of these were because of the rear seat exemption, and 2,485 (2.4 per cent) were because of the light truck and van exemption. In most states with seat belt laws (20), 4 to 8 per cent were exempted. However, in the six states that exempt light trucks and vans (all of which also exempt rear seat occupants), the per cent exempted ranged from 14 to 40. New Mexico (40 per cent), Oklahoma (33 per cent), and Missouri (25 per cent) have the highest exemption rates.

There is some limited evidence that exempted occupants do not increase their belt use to the same extent as occupants covered by the laws. Most state observational surveys of belt use have been restricted to front seat occupants and do not distinguish vehicle type. However, data from New Mexico show that seat belt use in passenger cars increased from 29 per cent before the law to 53 per cent early in the law's first year and was 46 per cent two years later. In contrast, belt use in pickup trucks in New Mexico was 19 per cent before the law, 24 per cent right after, and 18 per cent two years later.[5] In Texas, where both cars and pickup trucks are covered, belt use increased in both. Among car occupants, belt use was 16 per cent before the law, 64 per cent right after, and 56 per cent two years after the law. In Texas, 9 per cent of pickup truck occupants used seat belts before the law, compared to 49 per cent right after, and 43 per cent at the two-year mark.*

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*Unpublished data from Texas seat belt observations, 1987. Washington, DC, Insurance Institute for Highway Safety.

TABLE 1—Passenger Vehicle Occupant Deaths 1980–84 and the Extent to Which These Occupants Would Have Been Covered by Subsequent Seat Belt Use Laws

State	Total Deaths	Not Subsequently Covered: Rear Seat Occupants	Not Subsequently Covered: Light Truck, Van Occupants	Total Not Covered: N	Total Not Covered: %
California	12,005	0	0	0	0
Colorado	2,027	122	0	122	6
Connecticut	1,448	91	0	91	6
District of Columbia	110	7	0	7	6
Florida	7,328	505	0	505	7
Hawaii	430	24	0	24	6
Idaho	880	63	0	63	7
Illinois	5,079	300	0	300	6
Indiana	3,140	482	124	606	19
Iowa	1,619	85	0	85	5
Kansas	1,585	103	0	103	6
Louisiana	3,211	160	0	160	5
Maryland	2,021	105	232	337	17
Michigan	4,756	246	0	246	5
Minnesota	1,950	108	0	108	6
Missouri	3,171	175	619	794	25
Montana	887	0	0	0	0
Nevada	880	0	0	0	0
New Jersey	2,939	238	168	406	14
New Mexico	1,498	92	514	606	40
New York	6,085	419	0	419	7
North Carolina	4,365	347	0	347	8
Ohio	5,392	380	0	380	7
Oklahoma*	2,973	144	828	972	33
Pennsylvania	5,594	222	0	222	4
Tennessee	12,564	784	0	784	6
Texas	882	71	0	71	8
Virginia	2,955	195	0	195	7
Washington	2,272	0	0	0	0
Wisconsin	2,572	0	0	0	0
Total	105,631	5,579	2,485	8,064	7.6

*Oklahoma amended its seat belt law effective March 1, 1989 to include pickup trucks and vans.

Discussion

Seat belt use laws in the United States have been shown to be somewhat effective in increasing belt use and reducing injuries and fatalities.[6,7] However, their potential effectiveness is limited in part by excluding from coverage some vehicle occupants who have seat belts available for use. Because this study analyzed fatalities that occurred during the period when no seat belt laws were enacted, it does not measure the efficacy of seat belt use laws once in force, but it does illustrate that more of the population could have been included under the laws. Although most states have laws that would have applied to more than 90 per cent of the motor vehicle occupants killed in their states in the years just before the laws were passed, a few states—those that exempt light truck and van occupants but have many of these vehicles on their roads—exclude 25 to 40 per cent of their occupant population. The gaps in coverage may have arisen because of political exigencies in passing the laws, but they have no logical basis. Some of the gaps in coverage that existed in the original child restraint laws have been removed through amendments. Louisiana has amended its belt use law effective August 1, 1988 and Oklahoma amended its law effective March 1, 1989 to include vans and pickups. Similar amendments in other states would very likely increase the laws' effectiveness and save lives.

ACKNOWLEDGMENTS

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Johns Hopkins Injury Prevention Center Invites Nominations for 1989 Award of Excellence

In 1988 the Johns Hopkins Injury Prevention Center established the Award of Excellence to recognize and honor one or more individuals each year who have made substantial contributions to the prevention of injuries. Any person who has made significant achievements in the prevention of injuries, whether through research, policy formulation, advocacy, or other means is eligible for the award.

Garen J. Wintemute, MD, MPH, Assistant Professor of Family Practice at the University of California-Davis, was chosen as the first recipient in recognition of his work on the prevention of firearm injuries. The award was presented during a ceremony at the Johns Hopkins School of Hygiene and Public Health on October 26, 1988.

Nominations are invited for the second annual Award of Excellence, to be presented in the fall of 1989. Letters of nomination should include a description of the nominee's contributions to injury prevention and the significance of the nominee's achievements. A curriculum vitae of the nominee should accompany the nomination. Send nominations to:

Stephen P. Teret, JD, MPH
Director, Injury Prevention Center
Johns Hopkins University
615 N. Wolfe Street, Room 4027
Baltimore, MD 21205

Reducing Knee Morbidity among Carpetlayers

SHIRO TANAKA, MD, MS, SHIU T. LEE, MS, WILLIAM E. HALPERIN, MD, MPH, MICHAEL THUN, MD, MS, AND A. BLAIR SMITH, MD, MS

Abstract: Carpetlayers have a high prevalence of occupational knee morbidity, partly attributable to their use of the knee kicker to stretch carpet for wall-to-wall installation. While a mechanical alternative "power stretcher" is available, knee kickers are still widely used. A questionnaire survey indicated that unavailability of the mechanical stretcher at installation sites was a major factor for continued use of the knee kicker. Strategies to reduce use of the knee kicker are discussed. (*Am J Public Health* 1989; 79:334–335.)

Introduction

There are approximately 99,000 carpetlayers in the United States.[1] They file workers' compensation claims 108 times more than expected[2] for "knee joint inflammation attributed to kneeling, striking against a stationary object, etc." Both kneeling and use of the knee kicker were significantly related to their knee morbidity.[3]

A knee kicker is made of cast aluminum and weighs about 2 kg. Every carpetlayer owns one. To stretch carpet, it is pressed onto carpet by a hand and its padded end is kicked with the suprapatellar area. The average impact force exceeds 3,000 Newtons repeated about 140 times per hour[4] (Figure 1).

As an alternative, a mechanical stretcher using the lever and cogs is available. It weighs about 20 kg by itself; the total weight with accessories and the case may reach 40 kg (Figure 2). Customarily, it is provided by employers (contractor or carpet dealer) and assembled for on-site use. We examined the reasons for carpetlayers choosing between these two devices.

Method

As part of a health hazard evaluation on carpetlayer's knee problems, pertinent questions were included in a self-administered questionnaire to assess their use of the mechanical stretcher and subjective evaluation of the tool. The questionnaire was mailed to members of the local floorlayers union; only those who "have ever used a mechanical stretcher" were instructed to proceed with the subsequent questions.

Respondents were asked to estimate the ratio of using the two carpet stretching devices. Questions concerning the attributes of the mechanical stretcher included the ease of use, portability, safety to the knee, speed and quality of carpet installation, and availability at job sites. For each attribute, respondents were asked to choose one of four graded responses about the device with an implied comparison to the knee kicker.

From the Division of Surveillance, Hazard Evaluations and Field Studies, National Institute for Occupational Safety and Health, Cincinnati, Ohio. Address reprint requests to Shiro Tanaka, MD, MS, R-16, NIOSH, CDC, USPHS/DHHS, Robert Taft Laboratory, 4676 Columbia Parkway, Cincinnati, OH 45226. This paper, submitted to the Journal May 6, 1988, was revised and accepted for publication August 29, 1988.

FIGURE 1—A Carpetlayer Using a Knee Kicker

Results

One hundred and thirty-two (78 per cent) of 170 union members returned the questionnaire. The local union included subsets of other tradesmen, such as interior decorators, whom we excluded from our analysis, restricting it to the 94 (71 per cent) of respondents who reported ever using a mechanical stretcher. All respondents were males; 47 per cent used the knee kicker more often, 20 per cent used both tools about equally, and 33 per cent used the mechanical stretcher more often. Mean ages were similar in these three groups.

The majority of respondents reported that the mechanical device was easy to use, "saved" the knees, and achieved a better quality of installation more quickly (Table 1). However, almost 60 per cent said it was difficult to carry around; about one-third felt that it slowed down the job, and 39 per cent reported it was not always available at the job site.

FIGURE 2—A Carpetlayer Using a Mechanical Stretcher

TABLE 1—Favorable or Unfavorable Attributes of the Mechanical Stretcher (MS) with Implied Comparison to the Knee Kicker

Mechanical Stretcher (is):	Very Favorable About MS %*	Somewhat Favorable about MS %	Somewhat unfavorable about MS %	Very Unfavorable About MS %	Mechanical Stretcher (is):
Easy to use	47.9[a]	45.7[b]	6.4[c]	0[d]	Difficult to use
Easy to carry around	7.4	33.0	41.5	18.1	Difficult to carry around
Saves my knees	71.3	24.5	4.2	0	Harmful to my knees
Helpful in doing a quick job	33.0	34.0	29.8	3.2	Slows down my job
Helpful in doing a better job	87.2	9.6	3.2	0	Not as good as knee kicker
Available	60.7[e]	22.3[f]	17.0[g]	0[h]	Unavailable

*Percentages of each row add up to 100%.
a) Very easy, b) fairly easy, c) fairly difficult, or d) very difficult *to use.*
e) Always, f) usually, g) sometimes, or h) never *available.*

A logistic regression analysis showed that the "availability" and "save knee" factors were associated with the increased use of the mechanical stretcher.*

Discussion

Why do carpetlayers continue to use the knee kickers which are traumatic to their knees while a safer alternative is available?

Observation of carpet installation reveals that the knee kicker is kicked with two different intensities—mild and strong. Milder kicks are applied initially to engage the edge of carpet onto the tacks (which have been nailed to the perimeter of the floor), or when installing carpet in small confined spaces such as a closet.[4] In contrast, very forceful kicks are applied in normal installations to get a good stretch. The latter is likely to be most harmful to the knee, and should be totally replaced by using the mechanical stretcher. This survey sheds some light onto what could be done to increase the use of the mechanical device and reduce the use of the knee kicker.

The perceived shortcomings of the mechanical stretcher were: unavailability at the job sites, difficult portability (heavy and bulky), and cumbersome to use. The unavailability or non-use factor may be due to the initial higher cost of the mechanical stretcher ($400) compared to the knee kicker ($75) or failure of an employer to provide enough of the mechanical stretchers for each work crew to use; another economic factor might be that, if use of the mechanical stretcher slowed down the speed of installation for some carpetlayers, the employer would be reluctant to encourage its use.

*Availability: O.R. = 2.27, 95% CI = (1.35, 3.85); "Save" knee: O.R. 2.04, 95% CI = (1.02, 4.17); The questionnaire, detailed tabulations, and regression analysis data will be provided upon request to the first author.

Several strategies are suggested to increase the use of the mechanical stretcher:

- Building contractors, carpet dealers and trade unions could initiate or intensify educational efforts to emphasize the potential health hazard of using the knee kickers and encourage carpetlayers to rely more on the mechanical stretcher.
- Carpetlayers could be trained to become proficient in using the mechanical stretcher regardless of the room size.
- Employers should be encouraged to provide an adequate number of mechanical stretchers for each work crew.
- The portability issue could be addressed by such engineering approaches as attaching casters to the carrying case, or inventing a light-weight pneumatic stretching device.
- Consumers could be educated about the health hazards of the knee kicker and encouraged to specify the use of the mechanical stretcher for quality installation.
- A regulatory requirement could stipulate that employers provide each carpetlayer with a mechanical stretcher at job sites.

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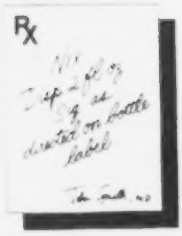

PEDICULICIDAL/OVICIDAL ACTIVITIES: *In vitro* data indicate that permethrin has pediculicidal and ovicidal activity against *Pediculus humanus* var. *capitis*. The high cure rate (97-99%) of Nix in patients with head lice demonstrated at 14 days following a single application is attributable to a combination of its pediculicidal and ovicidal activities and its residual persistence on the hair which may also prevent reinfestation.

INDICATIONS AND USAGE: Nix is indicated for the single-application treatment of infestation with *Pediculus humanus* var. *capitis* (the head louse) and its nits (eggs). Retreatment for recurrences is required in less than 1% of patients since the ovicidal activity may be supplemented by residual persistence in the hair. If live lice are observed after at least seven days following the initial application, a second application can be given.

CONTRAINDICATIONS: Nix is contraindicated in patients with known hypersensitivity to any of its components, to any synthetic pyrethroid or pyrethrin, or to chrysanthemums.

WARNING: If hypersensitivity to Nix occurs, discontinue use.

PRECAUTIONS:

General: Head lice infestation is often accompanied by pruritus, erythema, and edema. Treatment with Nix may temporarily exacerbate these conditions.

Information for Patients: Patients with head lice should be advised that itching, redness, or swelling of the scalp may occur after application of Nix. If irritation persists, they should consult their physician. Nix is not irritating to the eyes; however, patients should be advised to avoid contact with eyes during application and to flush with water immediately if Nix gets in the eyes. In order to prevent accidental ingestion by children, the remaining contents of Nix should be discarded after use.

Combing of nits following treatment with Nix is not necessary for effective treatment. However, patients may do so for cosmetic or other reasons. The nits are easily combed from the hair treated with Nix after drying.

Carcinogenesis, Mutagenesis, Impairment of Fertility: Six carcinogenicity bioassays were evaluated with permethrin, three each in rats and mice. No tumorigenicity was seen in the rat studies. However, species-specific increases in pulmonary adenomas, a common benign tumor of mice of high spontaneous background incidence, were seen in the three mouse studies. In one of these studies there was an increased incidence of pulmonary alveolar-cell carcinomas and benign liver adenomas only in female mice when permethrin was given in their food at a concentration of 5000 ppm. Mutagenicity assays, which give useful correlative data for interpreting results from carcinogenicity bioassays in rodents, were negative. Permethrin showed no evidence of mutagenic potential in a battery of *in vitro* and *in vivo* genetic toxicity studies. Permethrin did not have any adverse effect on reproductive function at a dose of 180 mg/kg/day orally in a three-generation rat study.

Pregnancy: *Teratogenic Effects*: Pregnancy Category B. Reproduction studies have been performed in mice, rats, and rabbits (200-400 mg/kg/day orally) and have revealed no evidence of impaired fertility or harm to the fetus due to permethrin. There are, however, no adequate and well-controlled studies in pregnant women. Because animal reproduction studies are not always predictive of human response, this drug should be used during pregnancy only if clearly needed.

Nursing Mothers: It is not known whether this drug is excreted in human milk. Because many drugs are excreted in human milk and because of the evidence for tumorigenic potential of permethrin in animal studies, consideration should be given to discontinuing nursing temporarily or withholding the drug while the mother is nursing.

Pediatric Use: Nix is safe and effective in children two years of age and older. Safety and effectiveness in children less than two years of age have not been established.

ADVERSE REACTIONS: The most frequent adverse reaction to Nix is pruritus. This is usually a consequence of head lice infestation itself, but may be temporarily aggravated following treatment with Nix. 5.9% of patients in clinical studies experienced mild temporary itching; 3.4% experienced mild transient burning/stinging, tingling, numbness, or scalp discomfort; and 2.1% experienced mild transient erythema, edema, or rash of the scalp.

DOSAGE AND ADMINISTRATION:

Adults and Children: Nix is intended for use after the hair has been washed with shampoo, rinsed with water and towel dried. Apply a sufficient volume of Nix to saturate the hair and scalp. Nix should remain on the hair for 10 minutes before being rinsed off with water. A single treatment is sufficient to eliminate head lice infestation. Combing of nits is not required for therapeutic efficacy, but may be done for cosmetic or other reasons.

SHAKE WELL BEFORE USING.

HOW SUPPLIED: Nix (Permethrin) 1% (wt./wt.) Creme Rinse is supplied in plastic squeeze bottles that contain 2 fl. oz. weighing 56 g. (NDC-0081-0780-81).

Store at 15°-25°C (59°-77°F).

1. Davies J, Dedhia H, Morgade C, et al: Lindane poisonings. *Arch Dermatol* 1983;119:142-144.
2. Taplin D, Meinking T, Castillero P, et al: Permethrin 1% creme rinse for the treatment of *pediculus humanus* var. *capitis* infestation. *Pediatr Dermatol* 1986;3:344-348.
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Health Effects of the War in Two Rural Communities in Nicaragua

Nicaragua Health Study Collaborative at Harvard, CIES,* and UNAN†

Abstract: We report on a pilot study to assess the effects of low intensity war in Nicaragua on the health of the civilian population. The study compared data from two regions in Nicaragua, one in an area of intense conflict, the other further removed from the war's violence. Information was obtained from a questionnaire administered to female heads of randomly selected households; structured interviews with community leaders and health workers; group discussions with community residents; and a review of regional and municipal death records. Height and mid-upper arm circumference of children were measured, and immunization records reviewed.

The war has had a serious negative effect on the lives of the civilian population in both the war zone and the non-war zone, with the effects most severe in the war zone. In both communities, over half of the respondents reported the death of a friend or relative. In the war zone community, over one-fourth of respondents reported attacks on family members in non-combat situations around their homes. Death by firearms was the leading cause of death in persons over age 6 in the war zone. Vaccination coverage, nutritional indices, and familial disruption were worse in the war zone community. The findings suggest that continued funding of the Nicaraguan *contra* forces by the United States may be harming the ostensible beneficiaries of that policy, and that use of such low intensity conflict as a foreign policy tool should be questioned. (*Am J Public Health* 1989; 79:424–429.)

Introduction

There has been a dramatic and well-documented improvement in the health status of the Nicaraguan population since the revolution of July 1979.[1–11] The war waged by counterrevolutionary forces (*contras*) attempting to overthrow the government of Nicaragua has threatened to halt or reverse these advances. The contra war in Nicaragua is an example of what is termed "low intensity conflict" (LIC) by United States military and diplomatic officials.[12–14] LIC refers to conflicts short of conventional war, which intentionally subject the civilian population to a combination of psychological, economic, and military pressure to promote the adoption of the desired social system by the civilian population.

Several reports, using national data and standard epidemiologic methods, have documented negative effects of the current war in Nicaragua on the health status of the civilian population.[10,15–17] To date, however, little field research has examined the effects of the war through the collection of data at the community level. We report here the results of a pilot study comparing health indicators collected in a uniform fashion in two regions in Nicaragua, one in an area of intense conflict, the other further removed from the violence of the war.

Methods

Data were collected on selected health indicators in two regions of Nicaragua: one suffering frequent *contra* attacks and another experiencing fewer contra attacks to serve as a comparison. For this pilot study, a single town was chosen in each region as representative of the region under study.

Information was obtained from four sources:

- A questionnaire administered to the female heads (*amas de casa*) of a random sample of households with at least one child under 6 years of age;
- Standardized interviews conducted with community leaders: health professionals, the mayor, school teachers, and the head of the women's organization;
- Facilitated group discussions;
- Review of municipal and regional death records.

Regional information was obtained in structured interviews with Ministry of Health officials and from annual health records.

The household questionnaire measured a variety of health indicators, including economic conditions, food availability, nutrition, child health, stability of family life, disruption or deaths due to the war, access to health services, educational status, and psychological health. Vaccination data were recorded on all children under 72 months of age whose vaccination cards were available. Standing height of all children of the household between the ages 24 and 71 months present during the interview was measured using a metal tape measure and right angle. Mid-upper arm circumference of the left arm of children between 12 and 71 months was measured using an insertion-type measuring tape.[18]

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Towns to be studied were selected from Nicaragua's Region V, an area in the central part of the country under increasing attack from the *contras* since 1984,[19] and from Region IV, an area on the Pacific Coast near the Costa Rican border. According to the Ministry of Health, Region IV has experienced only isolated *contra* attacks since the Revolution in 1979.

In Region V, we studied the town of Acoyapa. A community of about 8,500 population, and surrounded by nine small villages or *comarcas*, Acoyapa is located in the dry, hilly region of Chontales Province east of Lake Nicaragua, about 30 kilometers by paved and fairly secure road from the provincial capital, Juigalpa. While Acoyapa itself has not been subjected to a direct *contra* attack, the mayor reported there have been at least 10 attacks on the surrounding *comarcas*, and fighting has occurred within a few kilometers of the town.

El Ostional was selected as the comparison town from Region IV. A town of 441 residents, and also surrounded by neighboring *comarcas*, El Ostional is a fishing and farming community located on the southern Pacific coast of Nicaragua, 12 kilometers from the Costa Rican border and 30 kilometers by dirt road from San Juan del Sur, the nearest town with over 2,000 population.

Despite the difference in their sizes, Acoyapa and El Ostional are similar in several important aspects. The economies of both towns are based on food production, and both towns have health facilities with a permanent staff.

El Ostional and two towns in Region V of comparable size, economy, and proximity to a major road or city were pre-selected jointly by the Ministry of Health and the North American study participants. Arriving in the war zone in Region V, however, the team was informed by the regional Ministry of Health that one of the pre-selected towns had been attacked just 10 days earlier and that the road leading to the other town had been mined by the *contras*. Considerations of safety led to the final choice of Acoyapa on the day the study began.

The field work and data collection were conducted in January and February 1987 by 11 North Americans, five in the war zone and six in the non-war zone. Random samples of households were surveyed in both towns. The female head of household was defined as the woman within the household identifying herself to be the ama de casa. In Acoyapa, a group discussion was conducted with families of displaced persons. In El Ostional, all the women in the town were invited to participate in a group interview.

Results

Sample Survey

Fifty interviews were completed in 76 sampled households in Acoyapa. Thirty-nine interviews were completed in 59 sampled households in El Ostional. The selected households represented approximately 9 per cent of households in Acoyapa and 90 percent of households in El Ostional. Interviews were not conducted in sampled households if the female head of household was not home on three successive visits, or if no children under age 6 lived in the home. No interviews were refused in either community.

The two towns were similar economically, with nearly identical proportions of households having electricity (78 and 82 per cent in Acoyapa and El Ostional respectively), radio or television (50 and 51 per cent), and a latrine (84 and 82 per cent). Only half of the households in Acoyapa had piped water compared to 84 per cent in El Ostional. At the time of the investigation, however, the actual flow of water was disrupted in both communities due to malfunctioning pumping systems and parts being difficult to obtain because of the economic embargo. Houses in both communities were small and crowded, with a mean of about 4.5 occupants per room in both. A mean of roughly 2.5 children under 6 years of age per household was reported in both towns, although the mean number of occupants in Acoyapa was 10.7 compared to 7.8 in El Ostional.

Respondents were asked if their standard of living was better, worse, or the same as the year before. In Acoyapa (war zone), half (51 per cent) of the households reported their living standard currently *worse* than the year before compared to 41 per cent in El Ostional, while only 8 per cent in Acoyapa reported an *improved* living standard compared to 26 per cent in El Ostional (test for trend $\chi^2 = 3.13$, df = 1, p = .077).

Indices of change in food production, change in food availability, and nutrition were consistently worse in Acoyapa (war zone) than in El Ostional (non-war zone). Eighteen per cent (9 of 50) of the households in the war zone town had had one or more members leave agricultural production within the past year, compared to none (0 of 39) in El Ostional (95% CI = 1.7,∞).

As an indicator of availability of food, respondents were asked whether they were eating more, the same, or less of three basic starch foods and seven basic protein foods than the year before. A greater proportion of people in Acoyapa reported eating less rice, corn, chicken, milk, cheese, and fish, and fewer eggs and beans compared to the year before than in El Ostional (Table 1). Starch consumption had decreased in both communities, while protein consumption had decreased greatly only in Acoyapa (Figure 1). Consumption of both starch and protein foods had declined significantly more in the past year in Acoyapa than in El Ostional (Wilcoxon rank sum test: for starch foods $z = 3.04$, $p = .002$; for protein foods $z = 4.86$, $p < .0001$).

Prevalence of chronic undernutrition in children ages 24 to 71 months as estimated by the proportion of children with height for age below the 5th percentile* compared to the National Center for Health Statistics reference popula-

TABLE 1—Comparison of Change in Consumption of Basic Foods over Previous Year

Food	Number (per cent) of Households Eating Less of Food Item		Odds Ratio	95% CI
	Acoyapa	El Ostional		
Starches				
Corn	33 49 (67)	14 38 (37)	3.54	(1.45, 8.60)
Rice	38 49 (78)	22 38 (58)	2.51	(1.0, 6.37)
Plantains	21 48 (44)	11 38 (29)	1.39	(.58, 3.37)
Proteins				
Milk	27 49 (55)	7 38 (18)	5.44	(2.01, 14.7)
Eggs	21 49 (43)	5 38 (13)	4.95	(1.65, 14.8)
Beans	30 49 (61)	13 37 (35)	2.91	(1.20, 7.07)
Chicken	18 49 (37)	6 38 (16)	3.10	(1.09, 8.83)
Cheese	19 49 (39)	7 38 (18)	2.80	(1.03, 7.64)
Fish	11 47 (23)	2 36 (6)	5.19	(1.07, 25.2)
Red Meat	28 49 (57)	16 38 (42)	1.83	(.78, 4.32)

The analysis was done by comparing the proportion of households reporting eating less of a particular food to the proportion giving any of the other possible responses. Other possible responses were that the respondent eats more or the same of the food as the year before, or the respondent does not eat the food at all.

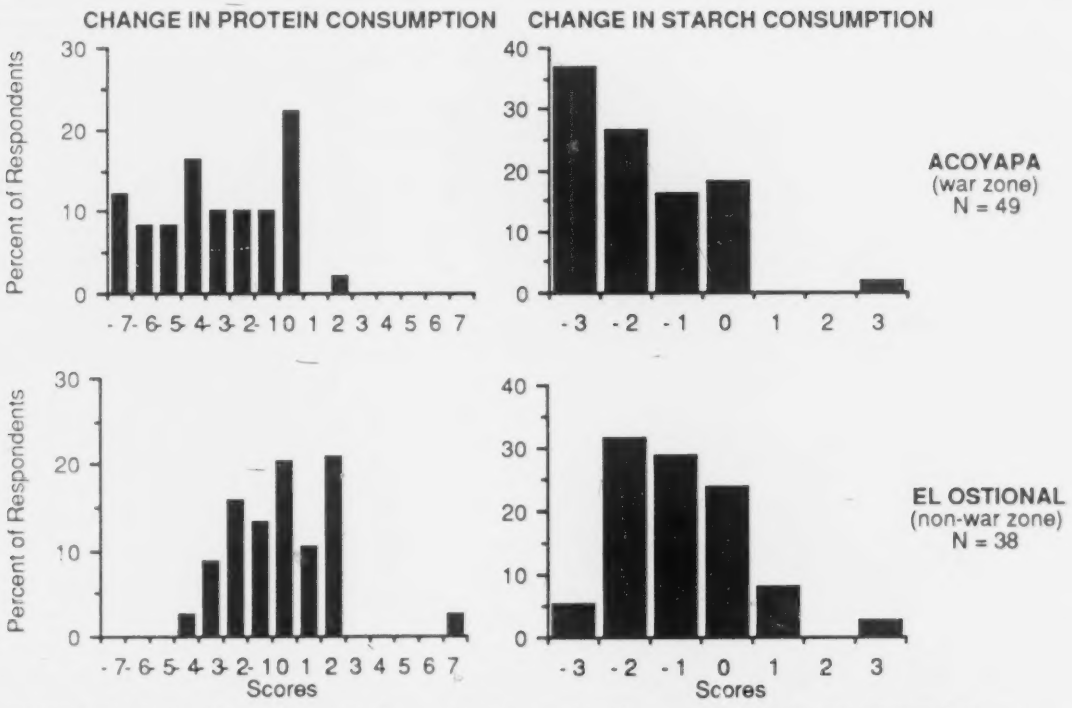

FIGURE 1—Distribution of households reporting increase, decrease, or no change in consumption level of protein and starch foods. To assess overall change in consumption of starch and protein foods, composite scores for change in both were constructed for each household. Integer values were assigned to responses for individual foods, and separately summed for the three starch foods and the seven protein foods. The response 'more' was scored with 1, 'less' with −1, and 'same' or 'not at all' with 0. Thus, starch scores range from −3 to 3 and protein scores from −7 to 7. Negative scores indicate a net decline in consumption of the seven protein and three starch foods investigated. Note the difference between towns in the proportion of households with protein scores −7 to −4 and with the worst starch score of −3.

tion[20-22] was greater in Acoyapa than in El Ostional. In Acoyapa (war zone), 22 of 50 children (44 per cent) were below the 5th percentile compared to 9 of 47 (19 per cent) in El Ostional (non-war zone) (OR = 3.32, 95% CI = 1.33, 8.29). Although many children in both towns showed signs of chronic undernutrition, none of the 81 children ages 12 to 71 months measured in Acoyapa and none of the 69 children measured in El Ostional exhibited signs of severe acute undernutrition as indicated by a mid-upper arm circumference less than 12.5 cm.[23-25]

Vaccination coverage was better in the non-war zone town than in the war zone town, based on both recorded completeness of vaccination and a crude measure of possession of an immunization card as a proxy for coverage.* In only 20 of 50 households (40 per cent) in Acoyapa compared to 29 of 39 (74.4 per cent) in El Ostional, immunization cards were presented for all children under 6 years of age in the household (OR = .23, 95% CI = .09, .57). Only 18 of 48 (37.5 per cent) children between the ages of 12 and 71 months with vaccination cards were completely vaccinated in Acoyapa compared to 47 of 60 (78 per cent) in El Ostional (OR = .17, 95% CI = .07, .39) (Table 2). A child was considered completely vaccinated if he or she had three recorded doses of DPT, three polio, 1 measles, and either a written record or scar for BCG.

Comparison of the immunization coverage shown in Table 2 with levels in 1980 shows that despite the war, considerable advances have been made. In 1980, the per cent of children one year old who were immunized was 50 per cent with BCG, 18 per cent with DPT, 18 per cent against polio, and 15 per cent against measles.[26]

Families in both Acoyapa and El Ostional have experienced violence as a result of the war. In both communities, over half of the respondents had experienced the death of a friend or relative, while one-fifth reported a member of the household killed, kidnapped, or wounded.

In Acoyapa (war zone), of the 26 respondents reporting that members of the family in the same household or relatives living in other households had been attacked or killed in the war, 14 indicated that the violence was done to non-

TABLE 2—Completeness of Vaccination Coverage in Children Ages 12–72 Months by Town

Vaccination	Number (Percentage) of children with complete vaccination*		Odds Ratio	95% CI
	Acoyapa n = 48	El Ostional n = 60		
All vaccinations complete	18 (38)	47 (78)	.17	(.07, .39)
3 DPT	19 (40)	56 (93)	.05	(.01, .15)
3 Polio	39 (48)	57 (95)	.23	(.06, .90)
1 Measles	35 (73)	51 (85)	.48	(.18, 1.28)
1 BCG	35 (73)	60 (100)	—	(−∞, .07)

*Sample excludes children who had no vaccination card

*Immunization cards are commonly kept in each child's house, in addition to records kept in the health centers.

combatant civilians in and around their homes. Of these, 13 were attributed to the *contras* and one was attributed to cross-fire between government forces and the *contras*. Two nieces of one respondent, both under 15 years of age, were kidnapped by the *contra* and still missing at the time of the interview. In one case, the *contra* murdered 11 male cousins and one of their wives at a family party, the respondent thought because of their cooperation with governmental agricultural extensionists. In another case, seven adults were wounded by the *contras*. In one incident, a man crossing lines of fire between the government forces and the *contra* was, according to the respondent, accidentally shot. No similar reports of violence against civilian relatives were offered in El Ostional.

The proportion of households in Acoyapa with children under 14 years of age missing at least one parent due to the war (17 per cent) was twice that observed in El Ostional (8 per cent) (OR = 2.32, 95% CI = .57, 9.47).

Mobility was high in both communities, with 16 per cent of the households in El Ostional and 21 per cent in Acoyapa reporting people moving in within the last year. Thirteen per cent of the households in El Ostional and 24 per cent in Acoyapa had people moving out within the past year. While none of the mobility in El Ostional was reported to be directly related to the war, half of the movement in Acoyapa was due to the war. A larger proportion of the *amas de casa* in Acoyapa (22.4 per cent) had lived in the town for less than three years than in El Ostional (8 per cent) (OR = 3.28, 95% CI = .84, 12.8).

Municipal Death Records

Review of the Municipal Registry of Acoyapa revealed that death by firearms was the leading cause of reported deaths in persons over 6 years of age in 1986, with a marked increase in death by firearms since 1984 (Figure 2). The municipality of El Ostional had no comparable registry. In Region IV, which includes El Ostional, records kept by the regional Ministry of Health showed death from firearms to be the fourth leading cause of reported deaths in persons over 6 years of age in both 1985 and 1986.

Regional Interviews and Records

The two regions studied differ greatly in the extent to which their primary health care systems have been affected by the *contra* war (Table 3). Ministry of Health records showed that in Region V (war zone), 20 health posts and health centers functioning in 1983 have been either destroyed, attacked, or closed due to the *contra* threat to health workers providing services there. By contrast, in Region IV (non-war zone), all of the clinics functioning in 1983 continue to provide services today.

The war has also affected programs of preventable disease control in Region V. Since 1984, five of ten health areas (administrative sub-divisions of the health region) in Region V have not been able to participate fully in the three campaigns held each year to administer measles, DPT, and polio vaccines to children. Half of the areas of Region V have also been unable to participate in the national malaria control campaign.

Community Interviews

In Acoyapa, civilians engaged in health and education work in the rural areas surrounding the town have been subject to attack by the *contra*. Health workers and teachers reported that at one health post, an auxiliary nurse was killed, and another salaried health worker threatened at gunpoint. Of

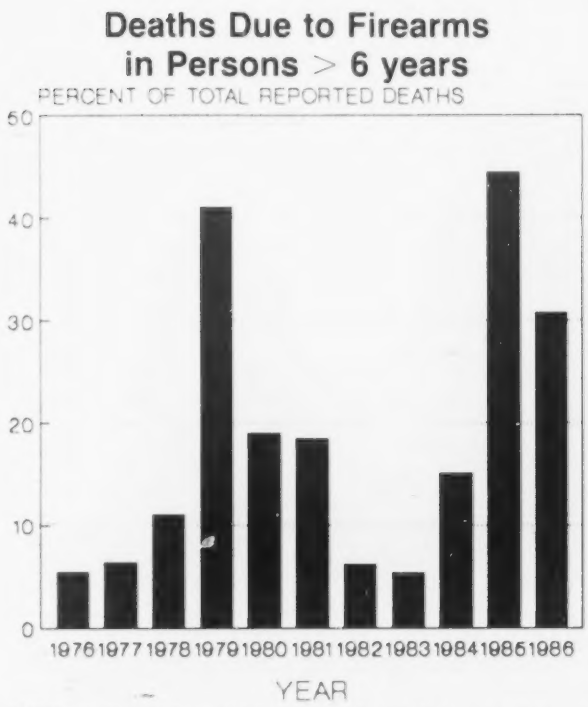

FIGURE 2—Percent of reported deaths due to firearms in persons over 6 years of age in the Municipality of Acoyapa (war zone) from 1976 to 1986. Absolute figures are 1976: 2/36; 1977: 2/31; 1978: 3/26; 1979: 14/34; 1980: 6/31; 1981: 7/38; 1982: 2/32; 1983: 2/38; 1984: 6/40; 1985: 24/54; 1986: 8/26. Source: Municipal Death Records Acoyapa.

TABLE 3—Regional Effects of the War

Health Service	Region V (war zone)	Region IV (non-war zone)
Health clinics:		
destroyed since 1984	3	0
attacked since 1984	2	0
closed since 1984	15	0
Health areas:		
not able to participate in vaccination campaigns	5/10	0
where malaria control workers could not provide services	5/10	0

SOURCE: Ministry of Health regional records

the 20 brigadistas (community health workers) serving the nine *comarcas*, ten have been kidnapped and five have quit due to the threat. Mobile clinics have been stopped and no vaccination campaigns occurred in 1986. Vaccination campaigns under the protection of army escort were planned for 1987. Four teachers in one of the *comarcas* near Acoyapa were threatened by the *contra*, and three requested transfers to other areas. The ongoing adult literacy campaign in the countryside has been discontinued due to this threat of violence.

The economic infrastructure of the area also has been affected by the war. According to the food distribution coordinator in Acoyapa, some 1,300 people moved into the town from its war-torn surroundings from 1983 to 1986. The market in Acoyapa where local producers used to sell their

food no longer takes place because people are too afraid to travel the roads frequently.

A group discussion with displaced persons from 20 different families in Acoyapa produced numerous reports of attacks on health and education facilities, as well as on civilians by the *contras*. These included the burning of an elementary school, an agricultural cooperative and two homes, the kidnapping of 14 people (including seven children), and the killing of at least ten people. An attack on a health center resulted in it, as well as the school, being closed.

The interviews with community leaders and the group discussion with women in El Ostional (non-war zone) showed that while that community has not been subject to armed conflict since the insurrection of 1978-79, the war has been felt in less direct ways. The area was shelled once from the sea in 1985. With the escalation of the war in 1983, the nearby border with Costa Rica was closed, ending the active trade that had existed between communities on either side of the border, and affecting availability of basic goods. Scarcity of resources has prevented the planned construction of a fish packing plant and child care center. Teachers reported young children in their classes expressing fear of violent death at the hands of the *contra*. Trenches have been dug around the health post to defend it from attack. The adult education supervisor reported there are areas in the region to which she no longer travels for fear of *contra* attack. In the group discussion with 35 women from El Ostional, mothers spoke extensively of their anxiety that their families would be harmed by the *contra*, or that they would not be able to adequately feed their families due to economic problems created by the war.

Discussion

The findings of this study provide a consistent picture of the damaging effects of the war on the civilian population in Nicaragua, especially in the war zone. Reports of deliberate targeting of health, education, and food production facilities by the *contras* are particularly apparent. Regional Ministry of Health officials in the war zone described a pattern of *contra* attacks on food production and transport facilities, as well as on the health care infrastructure. Community leaders and displaced persons interviewed in the war zone reported *contra* attacks on agricultural cooperatives, small farms, clinics, and individual health care and education workers in the countryside surrounding Acoyapa, the war-zone community. More than a fourth of the surveyed households in Acoyapa reported attacks on family members or relatives in non-combat situations in or around the families' homes. Attacks in the countryside are forcing people to leave subsistence farming, a principle source of food, and the population of Acoyapa has grown by 15 per cent in the past four years. Municipal death records show a marked increase in the firearm-related death rate among civilians in the Municipality of Acoyapa since 1984.

This study was a pilot of a method to assess the effects of low intensity conflict on the lives of a civilian population at the community level. The comparison of a community-based household survey in the war zone to one in the non-war zone was designed to measure the direct effects of *contra* or other combat activity on the civilian population. The method of random sampling was chosen to avoid possible bias in the respondents. Despite the relatively small number of households interviewed, a consistent pattern of more persons being forced from agriculture, worsening availability of food with attendant greater prevalence of chronic undernutrition, poorer vaccination coverage, a greater decline in the standard of living, as well as a greater direct experience of violence against civilians in the war zone emerges from the results of the household survey. The low prevalence of severe acute undernutrition in both towns suggests that existing food distribution systems, which include a government rationing system for staple commodities, have prevented frank starvation.

Several methodological problems limit interpretation of the comparative results of the sample survey. Generalizability of the results from the communities to their respective regions is limited because of the one-to-one comparison of the towns and the non-random selection of the communities to be studied. Differences in a one-to-one comparison may be due to many different factors, and particularly in this case to the difference in size of the towns studied. Information gathered from the other sources is, however, consistent with an interpretation that the differences are due to the war. Safety considerations prevented implementation of the original plan to study more than one town in the war zone and to study small towns of comparable size. In addition, it was impossible to blind the interviewers to the exposure status of the two study groups. The possibility of safety considerations and non-blinding of interviewers as sources of selection and interviewer bias, respectively, are two methodological problems inherent in studying the effects of low intensity conflict and political violence.[27]

Our findings are consistent with other reports by medical investigators from the US[28–31] that the *contras* have deliberately killed, kidnapped, and injured civilian health workers and destroyed health facilities. The *contras* have reportedly killed more than 100 Nicaraguan health workers.[32] Through May 1986, the Nicaraguan Ministry of Health reports that 65 health facilities have been completely or partially destroyed by the *contras*, and an additional 37 health units have had to be closed due to the threat of *contra* attack.[15] The disruption of vaccination and malaria campaigns has been associated with measles outbreaks and a rising prevalence of reported malaria cases in the areas of the country most affected by the war.[33] Of 4,429 war-related deaths reported in Nicaragua on the Nicaraguan side of the conflict through May 1986, one-third were civilian non-combatant deaths, including 210 children under 12 years of age.[15]

This study provides further documentation of the destructive effects the war against Nicaragua is having on the lives of civilians in two rural communities, with these effects being most severe in the war zone. This raises several important concerns for health professionals and the US public. The data suggest that continued US support for the *contra* war may lead to continued suffering of the Nicaraguan civilian population, the ostensible beneficiaries of US foreign policy in Nicaragua. Reports of the deliberate targeting of health and education facilities are particularly disturbing, raising questions of the violation of the Geneva Convention protection of health workers during war.[34] Lastly, the pattern suggested by this and other reports should lead the US public to consider carefully the effects of "low intensity conflict" on the welfare of civilian populations before its further application in the Third World.

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Call for Papers:
Society for Health and Human Values Annual Meeting

The Society for Health and Human Values has announced a call for papers for its annual meeting, to be held October 26–29, 1989 in Washington, DC. Topics are sought in the humanities and health sciences, including, but not limited to, ethics, philosophy, religious studies, literature, history, creative arts, and law. Social science, educational evaluation, and health policy studies related to the humanistic aspects of health care are also welcomed.

Papers should have a maximum reading time of 20 minutes. Provide author identification on a separate cover sheet to ensure blind review. *Send papers* by June 15, 1989 to: SHHV Annual Meeting Program, 6728 Old McLean Village Drive, McLean, VA 22101. For *meeting information*, contact the program chair, John Moskop, PhD, Department of Medical Humanities, East Carolina University School of Medicine, Greenville, NC 27858-4354. Tel: (919) 551-2797.

A Methodology for Targeting Hospital Cases for Quality of Care Record Reviews

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Abstract: We tested the efficacy of selected case characteristics in targeting quality of care problems for medical record review. The case characteristics, all of which apply to patients who die in a hospital, consist primarily of procedures and DRGs (diagnosis-related groups) for which death rarely occurs, and a set of complications of surgical care. All characteristics are obtainable from combinations of the principal and secondary diagnoses and procedures in the case, and are available from discharge abstracts.

The presence of a quality of care problem is confirmed through a review of the medical record by a nurse and two or more physicians.

A logistic regression model that controls for various patient and hospital variables is used as a measure of each of the proposed case characteristics. The results indicate that most of the characteristics are associated with higher percentages of quality of care problems than cases chosen at random, and that the methodology has promise as a tool for targeting cases for medical record review. (*Am J Public Health* 1989; 79:430–436.)

Introduction

In the last two decades, the costs of hospital care and other types of health care have risen more rapidly than the cost of living or the inflation rate. Partly as a result of the concentration of researchers and regulators on efforts to curtail these rising costs, hospital quality assurance initiatives have been relegated to a less important role than they deserve. Furthermore, it is feared that hospitals may have reduced diagnostic and therapeutic resources to patients, and discharged patients prematurely.[1]

Until recently, hospital quality assurance efforts have been quite limited. For the most part, surveys and quality of care reviews have relied on the first two of Donabedian's tripartite categorization of measures for judging quality: structure and process.[2] Quality of care researchers, regulators and review agencies have come to realize, however, that we must also incorporate Donabedian's third type of measure (outcome) into our reviews if we hope to accurately measure the quality of care in an institution.

A prime example of initiatives geared toward the use of outcome measures is the new agenda set forth by the Joint Commission on Accreditation of Healthcare Organizations, which is in the process of developing "clinical outcomes" to be reported by hospitals.

Several recent research studies have also used outcome measures to assess various aspects of quality of care in hospitals. A particularly well-researched issue has been the relation between the volume of patients treated at a hospital for a particular surgical procedure and the outcome achieved. Numerous studies reported that high surgical volumes for selected procedures are associated with better outcomes (lower mortality rates or shorter lengths of stay) for surgical patients.[3–6]

Another quality of care issue that has attracted considerable research attention is the "small area variations" in the rates of surgery for various procedures. Research by Wennberg and his associates[7–10] demonstrates that surgery rates among counties in the same state vary significantly even after adjusting for patient characteristics. This has led to the conclusion that physician practice patterns account for a significant percentage of observed differences.

Perhaps the most publicized efforts to develop outcome measures for quality of care in hospitals were the 1986 and 1987 reports by the Health Care Financing Administration (HCFA) that list hospitals with Medicare patient mortality rates significantly higher or lower than the national average.[11,*,†]

One shortcoming of the HCFA methodology involves the utility of the results for targeting purposes. Limited resources available to review organizations, state surveyors, accrediting agencies and hospitals as well as cost effectiveness concerns suggest a focus on reviews of selected *cases* (those most likely to evidence quality of care problems).

Another criticism of the HCFA model is that it has no proven link to the existence of quality of care problems. That is, it has not been proven that HCFA outlier hospitals have higher proportions of cases with quality of care problems (or deaths resulting from quality of care problems) than non-outlier hospitals. Recent studies by Dubois, *et al.*,[12,13] which use a methodology similar to HCFA's, are laudable in that a validation component is included. However, the database and the set of independent variables used in these studies are different from those used by HCFA.

The study described here is an attempt to identify hospital *cases* that are more likely to evidence quality of care problems than the average case. Our methodology is more suited than the HCFA methodology to integration with typical survey protocols, peer reviews, or hospital self-monitoring, all of which must identify specific hospital records to review.

The study has been confined to patients dying in a hospital. However, the number of in-hospital deaths generally exceeds the resources available for medical record reviews. Thus, the study attempts to define a subset of in-

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*Krakauer H: Prediction of statistical outliers. Health Care Financing Administration 1986, personal communication.

†Krakauer H: Outcomes in in-hospital care of Medicare patients in 1983 and 1984. Health Care Financing Administration 1986, personal communication.

hospital deaths for medical record review. It should be noted that the subset was not limited to deaths within 30 days of admission as in the HCFA study. The case characteristics that will be studied are referred to as targeting criteria. Although they apply primarily to surgical patients, some medical cases are included.

Methods

Identification of Targeting Criteria

The first step was to identify the case characteristics (targeting criteria) to be used in targeting a subset of in-hospital deaths in New York State in 1985 and 1986. To be effective as targeting criteria, these characteristics had to be retrievable from discharge data abstracts. Since 1980, every acute care hospital in New York State has been required to send a discharge data abstract to the State Department of Health for every inpatient stay. These abstracts contain information pertaining to demographics, dates of admission and discharge, diagnoses, procedures, admission status, and discharge disposition.

Given the available information and the decision to restrict the targeting to a subset of patients who died in the hospital, a set of 11 targeting criteria was proposed by a consultant surgeon. The intent was to select cases that would have a higher percentage of quality of care problems than cases without the criteria, as judged by medical record review. Appendix A describes each of the 11 criteria, which are not mutually exclusive.

There are two types of rare deaths; the only difference is that rare death diagnosis-related groups (DRGs) are used for the second criterion and rare death procedures are used for the first criterion. These rare death DRGs are similar to the "sentinel events" defined by Rutstein and his colleagues.[14]

Criteria (3) and (4), "death occurred within one day of surgery" and "death occurred two days after surgery," respectively, are based on the rationale that a patient dying shortly after surgery except for the exclusions noted may have experienced complications resulting from poor quality of surgical care or surgery that was inappropriate. The reason for using "within one day of surgery" and "exactly two days from surgery" was to evaluate any differences between the two. The intent was to combine these two criteria if both proved to be significant.

Criteria (5)–(9) include fluids/electrolyte imbalance, wound disruption or infection, acute renal failure, cardiopulmonary arrest, and other complications of surgery (mechanical complications of a device implant or graft, complications affecting specified body systems, and previously unspecified complications of surgical and medical care). For each of the criteria (5)–(9), the definition is such that the complication is coded as a secondary diagnosis in a surgical case with one or more non-diagnostic procedures. These cases are targeted because of the possibility that the complication occurred *during* the hospital stay.

Criteria (10) and (11), "burns from treatment" and "poisoning complications," are also specified as secondary diagnoses to increase the chance that they occurred during the hospital stay. They differ from criteria (5)–(9) in that they do not necessarily apply to surgical patients.

Case Sampling

To determine the effectiveness of the targeting criteria in identifying cases with quality of care problems, we identified retrospective samples of cases with and without the targeting criteria. We limited the sample selection to 104 hospitals reviewed by the Department's three downstate review agents. Resource constraints of the agents limited the total sample size to 8,109. Table 1 provides a breakdown of the sample by targeting criterion and review agent.

Since it was necessary to choose a sample with a sufficient number of cases having each of the targeting criteria, the sample was stratified so that it contained a higher percentage of cases with the targeting criteria than did the population of all in-hospital deaths. The resulting sample is not appropriate for estimating the percentage of quality of care problems that exist for cases with in-hospital deaths. For each case sampled, patient and facility characteristics (Appendix B) were also noted. These "control" variables were used in a multivariate model.

The first level of review of the medical record was performed by a registered nurse who determined whether the information appearing on the cover sheet (e.g., patient death, diagnoses, etc.) was correct. Any errors were noted and the correct information, determined after consultation with a medical record technician, was used to determine if the case should be reclassified as to the presence or absence of targeting criteria. Approximately 9 per cent of the cases

TABLE 1—Sample Sizes by Targeting Criterion and Review Agent

Targeting Criterion	IPRO**	CMR**	NYCHSRO†	Total	Percentage of Total Sample
Primary Surgical Procedures with Mortality Rate Less than 5%	330	141	135	607	7.5
DRGs with Mortality Rate Less than 5%	123	100	78	301	3.7
Death Occurred within 1 Day of any Surgical Procedure	675	378	317	1,371	16.9
Death Occurred on 2nd Day Following any Surgical Procedure	432	206	192	832	10.3
Surgical Case with Fluid Electrolyte Imbalance Reported as a Secondary Diagnosis	736	472	284	1,495	18.4
Surgical Case with Infection or Wound Disruption Reported as a Secondary Diagnosis	103	33	31	167	2.1
Surgical Case with Renal Failure Reported as a Secondary Diagnosis	55	30	70	155	1.9
Surgical Case with Cardiopulmonary Arrest Reported as a Secondary Diagnosis	173	75	224	472	5.8
Other Complications of Surgery	504	229	222	955	11.8
Burns Reported as a Secondary Diagnosis	10	14	6	30	.4
Poisoning is Reported as a Secondary Diagnosis	68	49	48	167	2.1
Total Unduplicated Targeted Samples*	2,833	1,518	1,213	5,564	68.6
Deaths not elsewhere classified	1,167	982	396	2,545	31.4
Total Samples	4,000	2,500	1,609	8,109	100.0

*Will be less than the sum of individual target groups because of cases with multiple target attributes.
**Island Peer Review Organization.
***Comprehensive Medical Review, Inc.
†New York County Health Services Review Organization.

contained errors requiring correction. The record was next examined for any evidence of quality of care issues via a subjective review. If the nurse had any doubts regarding the quality of care rendered, the record was referred to a physician for review. Each physician was a board-certified specialist or sub-specialist who had experience performing quality of care review and utilization review. Quality of care was judged subjectively by the physician in two ways: first, whether the care departed significantly from professionally recognized standards; and second, if there was such a departure, whether it caused or contributed to the patient's death. In those instances where the physician identified a quality issue, the hospital was notified and given the opportunity to respond. Upon receipt of the hospital's response, one or more physicians reviewed all the case material and issued a final determination. For simplicity of exposition, this step will hereafter be referred to as the second physician review.

It should be noted that physicians and nurses were not apprised of either the targeting criteria or the cases that had one or more targeting criteria.

Reliability

Because of resource constraints, reliability could not be measured accurately. Cases in which the first physician found no quality of care problem were not reviewed by a second physician. Cases in which the first physician found a quality of care problem were reviewed by one or more additional physicians, but the second reviews were conducted with additional information supplied by the hospital. Nevertheless, 66 per cent of the time the second physician agreed with the first physician that a problem existed. This concordance would probably have been higher if both physicians had reviewed the same record prior to additional input by the hospital.

Data Analysis

From the sampling and record review just described, the following information was available for each case (record):

- the presence or absence of a quality of care problem;
- which, if any, of the targeting criteria were present;
- the status of each control variable (age, sex, race, payor, county, admission status, hospital size, hospital sponsorship, hospital teaching status, utilization review (UR) agent, year of death, length of stay.

This information was used to determine if, for each of the 11 targeting criteria, the percentage of quality of care problems was higher than the corresponding percentage for cases without the characteristics.

Since we wanted to compare the proposed targeting criteria with other alternatives (such as patient and facility characteristics) for targeting, we used a logistic regression model to predict the presence of quality of care problems (the dependent variable) as a function of the targeting criteria and the various patient and facility characteristics (the control variables). Because the sample of cases was stratified to include an overrepresentation of cases with targeting criteria, a single logistic regression model would have underestimated the relative strength of each targeting criterion (because it would be contrasted with a sample having an overly rich representation of other targeting criteria). Thus, 11 separate regression models were formulated. Each model consisted of all cases with a given targeting criterion, and a simulated random sample. The random sample was chosen by picking targeted and non-targeted cases in proportion to their presence in the general population of hospital cases. No cases with the specific targeting criterion being tested were in the random sample, however. Thus, each regression model tested the ability of a specific targeting criterion to identify quality of care problems in comparison with a random sample of cases, while controlling for the factors presented in the Appendix.

Results

Table 2 presents the results of the targeting process by targeting criterion. Column (1) presents the number of cases selected by criterion; column (2) presents the number reviewed by a nurse. Differences between the numbers in the two columns represent cases for which medical records were not available or (rarely) for which a patient death was erroneously coded. The third column presents, by criterion, the number and percentage of cases reviewed by the nurse that were forwarded to the first physician. The percentages of cases forwarded ranged from 6.5 per cent for the non-targeted cases to 35.9 per cent for "other complications of surgery." The average percentage of targeted cases that were forwarded was 15.5 per cent, indicating that the nurse was about 2.4 times more likely to forward a targeted case than a non-targeted case; also, the nurse did not know which cases were targeted.

Columns (4) and (5) present the number and percentage of cases in which care departed from standards and in which care caused or contributed to patient death, respectively. The percentage of cases in which care departed from standards was 3.9 per cent for all targeted cases and 1.7 per cent for non-targeted cases. For targeted cases, the percentages ranged from 9.8 per cent for rare death DRGs to 2.2 per cent for fluid/electrolyte imbalance.

For quality of care problems that caused or contributed to patient deaths, the percentage was 2.4 per cent for all targeted cases and 1.0 per cent for non-targeted cases. For targeted cases, the percentages of quality of care problems ranged from 5.7 per cent for rare death DRGs to .6 per cent for poisoning complications.

Logistic Regression Results

The results of the logistic regression model for the quality of care definition "care that departed from professionally recognized standards" are presented in Table 3 for the targeting criteria. The odds ratio given for each criterion is an estimate of the relative chance of a case with the criterion having a quality of care problem compared to cases without the criterion, while holding other variables constant.

As is indicated in Table 3, all but three of the targeting criteria (fluid/electrolyte imbalance, burns and poisonings) proved to have significantly higher rates of quality of care problems independent of the control variables. Although cases having two or more criteria had a higher percentage of quality of care problems, no interactions between criteria were significant in the regression model. They comprised only 10 per cent of the cases. Also, no interactions between targeting criteria and control variables were significant.

Hospitalizations in New York County (Manhattan), review by New York County Health Services Review Organization, hospitalizations in proprietary facilities, patients with lengths of stay less than four days, hospitalizations in non-teaching hospitals, and Black patients were all associated with higher rates of quality of care problems.

Table 3 also presents analogous results for care that caused or contributed to patient deaths. In addition to the three targeting criteria that were not significant for the other

TABLE 2—Results of Record Review by Targeting Criterion: Total*

Targeting Criterion	(1) Total Cases	(2) Nurse Reviews	(3) No. of Cases Forwarded to First Physician (%) N	(3) No. of Cases Forwarded to First Physician (%) %	(4) Cases in Which Care Departed from Standards N	(4) Cases in Which Care Departed from Standards %	(5) Cases in Which Care Caused or Contributed to Patient Death N	(5) Cases in Which Care Caused or Contributed to Patient Death %
Primary Surgical Procedures with Mortality Rate Less than 5%	607	576	124	21.5	33	5.7	16	2.8
DRGs with Mortality Rate Less than 5%	301	264	65	24.6	26	9.8	15	5.7
Death Occurred within 1 Day of any Surgical Procedure	1,279	1,229	242	19.7	59	4.8	46	3.7
Death Occurred on 2nd Day Following any Surgical Procedure	832	802	141	17.6	35	4.4	20	2.5
Surgical Case with Fluids Electrolyte Imbalance Reported as a Secondary Procedure	1,419	1,384	128	9.2	30	2.2	17	1.2
Surgical Case with Infection or Wound Disruption Reported as a Secondary Diagnosis	167	161	29	18.0	7	4.3	3	1.9
Surgical Case with Renal Failure Reported as a Secondary Diagnosis	155	151	34	22.5	9	6.0	4	2.6
Surgical Case with Cardiopulmonary Arrest Reported as a Secondary Diagnosis	472	463	142	30.7	33	7.1	25	5.4
Other Complications of Surgery	992	955	343	35.9	35	3.7	24	2.5
Burn is Reported as a Secondary Diagnosis	30	27	3	11.1	2	7.4	1	3.7
Poisoning is Reported as a Secondary Diagnosis	167	162	12	7.4	4	2.5	1	6
Total of above	5,564	5,362	833	15.5	208	3.9	130	2.4
Total of non-targeted cases	2,545	2,479	161	6.5	43	1.7	24	1.0

*The total number of cases having each of the individual criteria is larger than the total because some cases have more than one criterion.

quality of care judgment, "wound disruption or infection" was also not significant. Thus, seven of the 11 criteria were significant for care that caused or contributed to patient deaths. Significant control variables were county (for four models), UR agent (two models), length of stay (one model), hospital sponsorship (four models), and primary payor (one model). For the variable "primary payor", Medicaid patients had higher rates of quality of care problems than patients of other payors.

Discussion

This study has demonstrated that information from uniform hospital discharge data sets can be used to target cases for quality of care review that have higher chances of having a quality of care problem than other cases. In contrast with methods for calculating adjusted hospital mortality rates for targeting, the approach presented here has been validated by medical record reviews. Furthermore, it is more useful in that it identifies specific cases in every hospital for review rather than just targeting a subset of hospitals. Also, certain

TABLE 3—Odds Ratios and Confidence Intervals for Targeting Criteria

Targeting Criterion	Care That Departed from Professionally Recognized Standards: Odds Ratio	Care That Departed from Professionally Recognized Standards: Confidence Interval for Odds Ratio	Care that Caused or Contributed to Patient Death: Odds Ratio	Care that Caused or Contributed to Patient Death: Confidence Interval for Odds Ratio
Primary Surgical Procedures with Mortality Rate Less than 5%	3.22	2.01, 5.15	3.26	1.76, 7.09
DRGs with Mortality Rate Less than 5%	5.19	3.20, 8.41	4.97	2.64, 9.35
Death Occurred within 1 Day of any Surgical Procedure	2.84	1.94, 4.17	3.54	2.21, 5.69
Death Occurred On 2nd Day Following any Surgical Procedure	1.99	1.58, 2.50	2.28	1.30, 3.98
Surgical Case with Fluid Electrolyte Imbalance Reported as a Secondary Diagnosis	1.13	.718, 1.79	1.17	.58, 1.76
Surgical Case with Infection or Wound Disruption Reported as a Secondary Diagnosis	3.04	1.07, 5.45	2.01	.73, 3.28
Surgical Case with Renal Failure Reported as a Secondary Diagnosis	3.18	1.52, 6.78	2.82	1.12, 7.09
Surgical Case with Cardiopulmonary Arrest Reported as a Secondary Diagnosis	3.40	2.08, 5.55	3.96	2.22, 7.06
Other Complications of Surgery	2.49	1.56, 3.97	2.41	1.40, 4.17
Burn is Reported as a Secondary Diagnosis	3.38	.75, 15.14	2.78	.71, 4.85
Poisoning is Reported as a Secondary Diagnosis	1.09	.38, 3.09	2.26	.25, 4.27

types of hospitals and patient characteristics are likely to be associated with quality of care problems.

Reweighting cases to simulate a random sample yields an overall percentage of 1.9 per cent for cases departing from standards and 1.2 per cent for cases in which care caused or contributed to patient deaths. Few studies in the literature provide a clue as to whether these percentages can be considered high or low. The study by Dubois, *et al.*[13] found more quality of care problems. Eighteen per cent of the deaths due to cerebrovascular accident, 24 per cent due to pneumonia, and 37 per cent of those due to myocardial infarction were judged to be "possibly preventable."

In a study conducted by the US Health and Human Services Department, 6.6 per cent of the 7,050 records reviewed were found to be receiving care that could be called poor quality—including improper diagnosis, treatment and therapy.[14] This percentage is about twice as high as that reported by the hospitals' federal peer review organizations.[15] The Medical Insurance Feasibility Study reviewed about 21,000 California hospital records and concluded that about 5 per cent of all admissions resulted in an injury caused by medical treatment.[16] A more recent study found that less than 1 per cent of hospital admissions are associated with poor care that results in temporary or permanent disability or death.[17] Thus, in general, the percentages reported in this study are lower than those reported in other studies. However, it is extremely difficult to compare results among studies because of the absence of uniformity in types of cases reviewed and in criteria for judging quality of care problems.

Despite the relative success of the study, several extensions would be beneficial. These include: enhancing the effectiveness of the targeting criteria, and developing other targeting criteria that will span the myriad of serious quality of care problems that occur in hospitals. Efforts to improve the effectiveness of the targeting criteria should focus on the incorporation of unambiguous severity of illness measures in definitions of the criteria. Severity of illness packages such as MedisGroups[18] and Computerized Severity Index[19] could be used to capture severity of illness at admission and then combined with the presence of post-admission complications to refine target definitions. The premise is that complications occurring during the hospital stay for patients with low severity of illness at admission will be more likely to be a result of poor hospital care.

With regard to the development of new criteria, two broad types of extensions can be explored. First, targeting can be expanded to include cases in which the patient did not die in the hospital. In some of these cases, the patient will have died a short time after discharge as a result of the complications. In other cases, the patient will not have died of the complications but will have experienced significant morbidity as a result of substandard hospital care.

The second means of expanding the current set of targeting criteria is to develop a set of medical complications to accompany the surgical complications. Some possibilities include infections; reactions from transfusions, injections and infusions; and vascular complications resulting from infusion, perfusion or transfusion. This extension is particularly important; in fact, many researchers feel that the majority of quality of care problems in hospitals are medical rather than surgical mishaps.

Another area that requires additional research is the development of protocols for quality of care determinations that are aimed at achieving maximum possible validity and reliability in decision making. As this was a pilot study, stringent validity and reliability tests were not included. However, given the relative success of the study, these issues will be addressed as part of a current effort of incorporating case targeting into New York State's new hospital surveillance program.

One particular concern is the validity of the judgments made by nurses in the initial screen. This is extremely important because most cases were not reviewed by physicians (they were screened out by nurses). In order to verify the validity of the nurse screens, a subset of cases screened out by nurses should be reviewed by physicians to ascertain that the physicians agree that no quality of care problems exist.

Unfortunately, limited resources did not allow for unbiased reliability tests in the study reported here. The best estimate, which was not a true estimate of reliability, was 66 per cent agreement on cases in which the first physician found a quality of care problem. However, we do not expect that any reliability problems would affect the *relative* percentages of quality of care problems for targeted and non-targeted cases, which was the focus of the study. Nevertheless, more work needs to be done in measuring the reliability of quality of care determinations and ensuring that it is at an acceptable level.

Our work should be interpreted in the light of related research. The study by Dubois, *et al.*,[13] is especially notable. As mentioned earlier, they used two types of quality of care review: a structured review that employed weighted process-of-care scales—11 for pneumonia and 10 each for cerebrovascular accident and myocardial infarction—and a review in which experts made subjective judgments as to whether the deaths were preventable. The former method found no differences in quality of care among high and low outlier hospitals whereas the latter method, which appears to be very similar to the subjective process used in this study, found that the high outliers had more preventable deaths than the low outliers. There is no way of determining whether either of the two results is valid. However, in assessing this anomaly, Dubois and colleagues imply that a major reason for the discrepancy may be that the structured approach was too myopic in only considering the specified criteria for noting a quality of care problem. With respect to reliability, however, the structured approach was superior.

The set of protocols and criteria developed by the Harvard Medical Practice Study Group in their ongoing study of medical injuries in New York State is also of value as a guide to improving validity and reliability.[‡] In this study, the Harvard Group defines an "adverse event" as a disabling injury inflicted upon a patient as a result of the process of medical management, whether it was negligent or not. Then, for adverse events, the probability of causation (how probable the condition was attributable to medical management rather than to the disease) and the probability of negligence (the probable contribution of negligence in medical management to the adverse event) are estimated.

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APPENDIX A
Definition of Mortality Study Targeting Criteria

Criterion	Description
1) Primary Surgical Procedures with Mortality Rate Less Than 5%	Principal procedure associated with mortality less than 0.5 per cent in 1985 Exclusions: None
2) DRGs with Mortality Rate Less than 5%	DRGs associated with mortality less than 0.5 per cent in 1985 Exclusions: None
3) Death Occurred within 1 day of any Surgical Procedure	Death occurred on same date or day following surgical date Exclusions: • Patients with any diagnosis of selected neoplasms (196.0 to 198.89)* • Patients with unscheduled admission and admitting or primary diagnosis of injury or poisoning (800.00 to 995.89) • Selected surgical procedures • Diagnostic and therapeutic principal procedures (87.0 to 99.99)
4) Death Occurred on 2nd Day Following any Surgical Procedure	Death occurred on second day following surgical date Exclusions: Same as for Criterion 3
5) Surgical Case with Fluid Electrolyte Imbalance Reported as a Secondary Diagnosis	Surgical case with a secondary diagnosis of fluid electrolyte, or acid-base balance (276.0 to 276.9) Exclusions: Same as for Criterion 3, and • Admitting or primary diagnosis of fluid electrolyte, or acid-base balance (276.0 to 276.9)
6) Surgical Case with Infection or Wound Disruption Reported as a Secondary Diagnosis	Surgical case with a secondary diagnosis of wound disruption (998.3) Exclusions: • Diagnostic and therapeutic principal procedures (87.0 to 99.99)
7) Surgical Case with Renal Failure Reported as a Secondary Diagnosis	Surgical case with a secondary diagnosis of urinary complication (997.5) Exclusions: • Diagnostic and therapeutic principal procedures (87.0 to 99.99)
8) Surgical Case with Cardiopulmonary Arrest Reported as a Secondary Diagnosis	Surgical case with a secondary diagnosis of cardiac complication (997.1) Exclusions: • Diagnostic and therapeutic principal procedures (87.0 to 99.99)
9) Other Complications of Surgery	Surgical case with a secondary diagnosis of mechanical or other body system reaction (996.00 to 999.9) Exclusions: • Diagnostic and therapeutic principal procedures (87.0 to 99.99) • Patients with no secondary procedure • Cases with Criteria 7) and 8)
10) Burn is Reported as a Secondary Diagnosis	Burns as a secondary diagnosis (940.0 to 949.5) Exclusions: • Admitting or primary diagnosis of burn (940.0 to 949.5)
11) Poisoning is Reported as a Secondary Diagnosis	A secondary diagnosis of selected drug poisonings (960.0 to 989.0) Exclusions: • Admitting or primary diagnosis of selected drug poisonings (960.0 to 989.9)

*Entries in parentheses are ICD-9-CM codes.

APPENDIX B
Control Variables Used in the Study

Variables	Description
Sex*	Male
	Female
Race*	White
	Black
	Hispanic
	Other
Age*	Continuous
Payor*	Blue Cross
	Medicare
	Medicaid, Other Government
	Other Commercial
	Other (Workers Comp., No Charge, Self-Pay)
County	Nassau
	Suffolk
	Bronx
	Kings
	New York
	Queens
	Richmond
Admission Status*	Unscheduled
	Scheduled
Hospital Size	Small (1–275)
	Medium (276–900)
	Large (901+)
Hospital Sponsorship	Proprietary
	Voluntary, Public
Hospital Teaching Status	No
	Yes
Utilization Review Agent	Island Peer Review Organization
	Comprehensive Medical Review Inc.
	New York County Health Services Review Organization
Year of Death*	1985
	1986
Length of Stay*	1–3 days
	4–14 days
	15–30 days
	31–60 days
	>60 days

*Contained in the Statewide Planning and Research Cooperative System (SPARCS)

APHA Releases Report on 'Illicit Drug Use and HIV Infection'

Illicit Drug Use and HIV Infection—The 3rd report of the Special Initiative on AIDS, a program activity of the American Public Health Association—has been released. This new report, prepared by the AIDS Working Group, summarizes the characteristics of illicit drug use and the scope and distribution of the HIV epidemic among intravenous (IV) drug users. Strategies to reduce the risk of HIV infection among drugs users are reviewed, including the need for education, health care, and drug treatment services upon demand, and strategies to promote safe sex.

To date, three reports have been issued in this AIDS Report Series. No. 1, *Casual Contact and the Risk of HIV Infection* was released July 1988; No. 2, *Contact Tracing and Partner Notification*, November 1988; and *Illicit Drug Use and HIV Infection*, January 1989. Single copies of the various reports are available for $3 each prepaid from APHA Publication Sales, 1015 15th Street, NW, Washington, DC 20005. Bulk rates available on request.

Coronary Arteriography and Coronary Bypass Survey among Whites and Other Racial Groups Relative to Hospital-based Incidence Rates for Coronary Artery Disease: Findings from NHDS

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Abstract: To assess racial differences in health care utilization for coronary artery disease (CAD) the data of the National Hospital Discharge Survey (NHDS) from 1979-84 were examined. Discharge rates for acute myocardial infarction (AMI) were utilized as a measure of hospital-based incidence and relative need for the designated cardiac procedures. Although 35-74 year old Black men had discharge rates of AMI that were 77 per cent of those observed for White men, they underwent coronary arteriography half as often and were only a third as likely to have coronary artery bypass graft (CABG) surgery. Black women in this age range were hospitalized at a slightly higher rate than White women for AMI, yet experienced a 19 per cent lower rate of coronary arteriography and a 52 per cent lower rate of CABG surgery. These data suggest a racial bias in the pattern of care delivered for CAD in US hospitals at the present time. (*Am J Public Health* 1989; 79:437–440.)

Introduction

Despite a continuing controversy about the real rates of both incidence and mortality among Blacks from coronary artery disease (CAD), it remains by far the most common cause of death for both Blacks and Whites in the United States.[1,2] A higher rate of CAD at a younger age among Blacks, with a cross-over of age-specific rates around 70, is observed in the vital statistics data,[3] and this finding has recently been replicated in hospital-based incidence rates.[4] Several reports have also demonstrated shorter average survival among symptomatic Black patients with CAD compared to Whites,[5,6] especially among women.[6] Although coronary artery bypass graft (CABG) surgery probably has a minor effect—if any—on survival rates from CAD in the population, the quality of secondary prevention and post-MI care overall could be important.[7] Previous reports suggest that Blacks undergo fewer cardiac diagnostic procedures and subsequent surgery than do Whites.[8–10] We further examined the issue of racial differentials in the diagnosis and treatment of CAD in the random sample of US hospitals provided through the National Hospital Discharge Survey.

Methods

National Hospital Discharge Survey (NHDS) data tapes for the years 1979-84 were obtained from the National Center for Health Statistics (NCHS). The NHDS is an ongoing sample survey drawn from a frame of all short stay non-federal hospitals in the US. A description of the design and procedures of the NHDS has been published.[11]

Data from these tapes were analyzed for coronary arteriography (ICD9 CM codes 88.55 - 88.57), CABG surgery (ICD9 CM codes 36.10 – 36.19) and acute myocardial infarction (AMI) (ICD9 CM code 410). Race, which is not always found on discharge summaries, was missing from 8.6 to 13 per cent of records from 1979-84. Persons with race not stated were excluded from the present study. Procedures for identifying the race category "other" varied from year to year but consisted primarily of aggregating Blacks, native Americans, and Asians. For 1978, we determined that approximately 88 per cent of the recoded race category "other" was Black. Therefore, the use of the term Black must be regarded with caution, and it should be realized that some individuals of other races are included in this category.

Age, race, and sex specific population estimates of the numbers of AMI and cardiac procedures were obtained. We limited our analysis to the ages 35-74 years (grouped into 10 year age intervals) because outside this age range the numbers of records contained in the NHDS tapes for Blacks are often very small, and the number of procedures and AMI between the ages of 35 and 74 years constitute about 95 per cent of the total when all ages are considered. Population-based rates were calculated using the age, race (White and Black), and sex-specific estimated US civilian population data as the denominator. For the year 1980 the actual census data were used. Race-sex specific hospital discharge rates were age-adjusted to the 35-74 year old US population of 1980. Relative standard errors (RSEs) were derived from relative standard error curves published annually by the NCHS and documentation provided by the NCHS.* All standard errors are expressed as approximate relative standard errors (RSE). Separate examination of age-specific rates did not demonstrate a pattern that varied from the age-adjusted rates.

Hospitalization rates of AMI were utilized as an index of relative need for health care for CAD. In a previous analysis of NHDS data from 1972 through 1984 our group demonstrated that age-adjusted in-hospital case fatality rates from AMI were 19 per cent for Whites and 18 per cent for Blacks,[4] thus ensuring that the diagnosis, as reflected in the severity of outcome, was similarly applied in the two race groups. As estimates of health care utilization relative to need, ratios were generated between rates of arteriography, CABG surgery and AMI in the sex-race groups. Finally, the relation of principal source of payment and type of hospital to the number of discharges for AMI and procedures performed for

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*I. Shimizu, PhD, NCHS, personal communication.

TABLE 1—Rate of Coronary Arteriography, Coronary Artery Bypass Graft Surgery and Acute Myocardial Infarction per Year in the US for the Period 1979–84, Persons Age 35–74, by Sex and Race, The National Hospital Discharge Survey (per 100,000, Mean ± RSE)

	Black Men		White Men				Black Women		White Women			
	Rate	RSE (%)	Rate	RSE (%)	B:W* Ratio	RSE (%)	Rate	RSE (%)	Rate	RSE (%)	B:W Ratio	RSE (%)
Coronary Arteriography	176	9.6	330	2.7	0.53	10.0	107	5.9	132	3.2	0.81	6.7
Coronary Artery Bypass Surgery	103	6.3	294	2.8	0.35	6.9	34	7.9	71	3.8	0.48	8.2
Acute Myocardial Infarction	641	3.4	833	1.7	0.77	3.8	338	4.1	328	2.2	1.03	4.4

*Black:White

White and Black patients was examined using data collected by the NHDS.

The per cent of records with race not stated, the relatively small number of Blacks and other minorities sampled, and other methodological problems which arise with the use of the NHDS data for epidemiologic research[12] suggest caution in interpreting the data to be reported. Despite these reservations, the NHDS is the only data set allowing estimates of population rates of the procedures of interest, and the large proportion of Blacks in the "other" category permits at least tentative conclusions. Data analysis was performed on an IBM 3081K64 computer utilizing programs available on SAS.[13]

Results

Average rates for coronary arteriography, CABG surgery, and AMI are summarized for the six years in Table 1. Black men, when compared to White men, were only half as likely to undergo arteriography and a third as likely to have CABG surgery performed, despite an AMI rate of 77 per cent of that of White men. Less differential was observed in relative terms for women, with an arteriography rate of 81 per cent in Blacks compared to Whites, and CABG surgery rates one-half that of Whites.

In an effort to provide a summary index of health care utilization relative to need, we generated ratios of the rates of the two procedures to the rate of AMI for the four sex-race groups. For every 2.5 White men (RSE 3.2 per cent) and 2.5 White women (RSE 3.9 per cent) discharged with the diagnosis of AMI, one arteriogram was performed, while only one in 3.6 Black men (RSE 10.2 per cent) and one in 3.2 Black women with an AMI (RSE 7.2 per cent) had this procedure. For every 2.8 White men (RSE 3.3 per cent) discharged with AMI a CABG procedure was recorded, with the comparable figure of 6.2 (RSE 7.2 per cent) to one for Black men. Corresponding ratios of 9.9 (RSE 8.9 per cent) and 4.6 AMI per CABG surgery (RSE 4.4 per cent) were recorded for Black and White women, respectively. Thus, if the assumption is made that the ratio of AMI to arteriography and CABG surgery observed among Whites is the norm in the community, these secondary ratios provide an estimate of the degree to which Blacks are underserved. Despite varying rates of individual procedures/diagnoses, the estimate of health care utilization deficit is remarkably stable in the two sexes, yielding a 31 per cent (RSE 10.7 per cent) deficit among Black men and 22 per cent (RSE 8.2 per cent) deficit among Black women for arteriography relative to rates of AMI among Whites and a 50 per cent (RSE 7.9 per cent for men and 9.9 per cent for women) lower rate of CABG surgery for both sexes among Blacks relative to Whites.

The per cent distribution of White and Black patients by hospital type ownership is presented in Table 2. Although the data suggest some differences in the type of hospitals where care is delivered, the large standard errors preclude any definitive conclusions. However, Whites are less frequently hospitalized in proprietary hospitals for CABG surgery. Concerning sources of payment, reimbursement for AMI for Blacks was more likely to originate from various government programs whereas reimbursement for coronary arteriography was more likely to originate from Medicaid in addition to various government programs (Table 3).

TABLE 2—Per Cent Discharges ± One Approximate Relative Standard Error in Per Cent for Acute Myocardial Infarction, Coronary Arteriography, and Coronary Artery Bypass Surgery According to Hospital Ownership, Combined National Hospital Discharge Survey Data from 1979–84

	Acute Myocardial Infarction		Coronary Arteriography		Coronary Artery Bypass Surgery	
	Blacks	Whites	Blacks	Whites	Blacks	Whites
Weighted N	290,547	3,713,220	77,808	1,113,450	32,763	854,063
Proprietary	10.68 ± 11.9	8.00 ± 6.2	6.57 ± 21.8	5.42 ± 12.0	17.05 ± 21.2	5.93 ± 12.5
Government	25.26 ± 9.0	19.38 ± 5.2	13.77 ± 23.0	9.17 ± 10.9	13.94 ± 22.3*	11.05 ± 11.0
Nonprofit:						
Church	18.75 ± 10.0	19.67 ± 5.2	23.18 ± 15.8	23.22 ± 9.0	16.59 ± 21.3	23.13 ± 9.5
Other	45.13 ± 7.8	52.95 ± 4.8	56.48 ± 12.9	62.19 ± 9.9	52.41 ± 16.2	59.89 ± 8.0

*Unstable estimate

TABLE 3—Per Cent Distribution of Acute Myocardial Infarction, Coronary Arteriography, and Coronary Artery Bypass Graft Surgery ± One Approximate Relative Standard Error in Per Cent According to Principal Source of Payment, Combined National Hospital Discharge Survey Data from 1979–84

	Acute Myocardial Infarction		Coronary Arteriography		Coronary Artery Bypass Graft Surgery	
	Blacks	Whites	Blacks	Whites	Blacks	Whites
Weighted N	290,547	3,713,220	77,808	1,113,450	32,763	854,063
Various government†	5.30 ± 14.3	2.89 ± 7.2	8.19 ± 20.4	3.03 ± 13.8	3.70 ± 33.7*	2.86 ± 14.8
Medicare	36.22 ± 8.1	44.00 ± 4.9	18.09 ± 16.8	24.46 ± 9.0	25.50 ± 19.0	28.44 ± 9.1
Medicaid	21.38 ± 9.5	18.35 ± 5.2	14.42 ± 17.8	7.16 ± 11.4	14.06 ± 22.2*	9.02 ± 11.5
Private insurance	31.40 ± 8.6	31.95 ± 5.0	54.62 ± 213.0	62.66 ± 7.7	48.77 ± 16.3	57.74 ± 8.0
Miscellaneous‡	5.69 ± 14.2	2.81 ± 4.9	4.69 ± 23.9*	2.64 ± 14.1	7.98 ± 26.0*	1.95 ± 16.2

*Unstable estimate
†Includes workers' compensation, Title V, and other government programs
‡Includes self-pay, no charge, and other sources of payment

Discussion

The results reported here suggest that, between 1979 and 1984, the Black population in the US received substantially less tertiary health care for CAD as measured by coronary arteriography and CABG surgery than Whites relative to a hospital-based measure of AMI incidence. It should be clear that these data do not reflect true incidence rates of CAD in the community; in fact, higher out-of-hospital sudden CAD rates have been reported among Blacks, and there is at least some evidence that CAD is diagnosed late at the catheterization laboratory in some centers.[1,2,14] True population rates of CAD, therefore, are likely to be even higher among Blacks, particularly women. This study examined only the response of US hospitals to patients that were discharged with one form of CAD, namely AMI. As noted previously the in-hospital case fatality rates from Blacks with AMI was identical to that of Whites,[4] suggesting at least that there was no bias toward milder cases. Data from at least one study indicate that the rate of normal coronary arteries found at catheterization is much higher in Blacks than Whites, and this would reduce the need for subsequent CABG surgery in that group.[15] In the context of the NHDS data for AMI it seems most likely, however, for the US as a whole this reflects the fact that appropriate patients are not receiving catheterization, rather than a true low rate of CAD.

Tertiary care for CAD is expensive and this fact helps to explain the race differences in health care utilization observed in this study. This study presents evidence that various government programs shoulder a greater burden for payment for hospitalizations for AMI and cardiac procedures among Blacks. In an entirely separate national probability sample, Young noted that the rate of CABG surgery in privately insured or self-paying patients was 9 per cent or higher, and only 2 per cent for those on public aid.[10] Multivariate analysis confirmed the importance of economic factors, and showed higher rates among Whites; higher CABG surgery rates were likewise observed in more affluent areas with more physicians and hospital beds relative to the population.[10] At a single university hospital it was observed that the rate of CABG surgery following arteriography was much lower among Blacks than Whites, and this could not be explained by clinical characteristics of the patients.[8] Maynard, *et al.* reported the same pattern from the data of the Coronary Artery Surgery Study.[9] Evidence suggests that the level of care for cardiovascular disease for Blacks is lower than for Whites in other measures of health care utilization.[16,17]

The present analysis suggests that these differentials reflect a biased institutional response since the burden of CAD reported in the same survey does not justify lower rates of cardiac procedures in Blacks. In view of increasing pressure for cost containment, and the growing reluctance of private hospitals to serve patients without adequate financial resources, it seems unlikely that these differentials will disappear in the foreseeable future. Recent evidence that the decline in CAD over the past 10 years among Blacks is slower than that of White men[18] further suggests that the unmet need for CABG surgery will remain greater among Blacks than among Whites.

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Improving the Homogeneity of Diagnosis-Related Groups (DRGs) by Using Clinical Laboratory, Demographic, and Discharge Data

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Abstract: For 48 of the most common diagnosis-related groups (DRGs) at our hospital, we examined the ability of clinical laboratory tests, demographic data, and ICD-9-CM codes, which provide a measure of severity of illness, to predict patients' length of stay (LOS) more accurately than DRGs alone. For 10 of 20 medical DRGs and 13 of 23 surgical DRGs examined, we were able to increase the ability to predict LOS by at least 10 per cent. The laboratory tests that proved most predictive of LOS over all DRGs were the mean serum sodium, potassium, bicarbonate, and albumin. The system is data driven, objective, and flexible, thus ensuring its utility for the purpose of equitable reimbursement. (*Am J Public Health* 1989; 79: 441–444.)

Introduction

Diagnosis-related groups (DRGs) are currently used to determine reimbursement for Medicare inpatients under the recently instituted prospective payment system (PPS). DRGs, however, may not accurately reflect an individual patient's requirements for hospital resources, due, in part, to the absence of a severity-of-illness measure.[1]

In many institutions, computerized clinical laboratory systems allow access to a patient's test values in machine-readable form. In this study, we provide preliminary evidence that patient data from these sources can improve the ability of DRGs to predict a patient's hospital resource utilization, as proxied by length of stay (LOS). The modified DRGs are produced by an automated, objective, and data driven procedure, eliminating the need for potentially biased or laborious chart review.

Methods

The source of patient data for this study is the Combined Patient Experience (COPE) database, created by merging patient data from computerized sources at the University of California, San Francisco (UCSF) Medical Center.[2] The data include clinical laboratory test results, hospital admission and demographic data, and International Classification of Diseases, 9th Revision, Clinical Modification (ICD-9-CM) diagnosis and procedure codes from discharge abstracts. COPE data tapes from the first nine months of 1981 and 1982 were used, including 15,910 and 15,930 admissions, respectively. (The first nine months of data were used because data from the fourth quarter of 1982 were not readily available as a merged data set.) All patients were classified into one of 470 DRGs using the Health Systems International DRG Grouper, running on a Digital Equipment Corporation MicroVAX II supermicrocomputer.

Selection of a DRG for analysis was based on the frequency of admission to the hospital of patients in that DRG, with the goal of examining the top 10 per cent of DRGs. Forty-eight DRGs were chosen, 28 surgical and 20 medical, encompassing 12.7 per cent of all medical-DRG patients (2,027 patient) and 16.7 per cent of all surgical-DRG patients (2,650 patients) for the 1981 time period; figures for 1982 were comparable. Neonatal and obstetric/gynecologic DRGs were excluded from analysis because of a paucity of clinical laboratory test results for patients with those DRGs. Patients whose LOS made them outliers, as defined by Medicare for the DRG to which they were assigned, were excluded from analysis.

The patient attributes selected as candidates for modifying DRGs were age, sex, laboratory test results, ICD-9-CM discharge and diagnosis codes, and status (alive or dead) at discharge. The laboratory tests used were components of the complete blood count (hemoglobin, white blood count); and commonly ordered serum chemistries (sodium, potassium, chloride, total carbon dioxide, blood urea nitrogen, creatinine, albumin, glucose, total protein, calcium, cholesterol, uric acid, total bilirubin, alkaline phosphatase, lactic dehydrogenase, glutamic oxaloacetic transaminase, and phosphorus).

The measures of resource utilization available in the COPE database were LOS and total charges; these were found to be highly correlated with one another ($r = 0.9$), and LOS was chosen as the outcome variable.

All test results were log-transformed (to decrease skewness) and were then centered to the log of the midnormal value for that test and scaled to unit variance. Results were then trimmed at the 5th and 95th percentile values (to delete outliers) by replacing values above or below those percentiles with the values precisely at those percentiles. For each patient admission, the time-weighted mean value for each (transformed) laboratory test was computed by taking the area under the curve of the laboratory test value against time of hospital stay (using the trapezoidal rule), divided by the LOS. In so doing, the test-versus-time curve was extended horizontally from the first (or last) test value to the time of admission (or discharge), if test values on those days were unavailable. When a selected test was not present in a case record, the result was assigned the mid-normal value of that laboratory test.

In addition to the laboratory test variables, the presence or absence in the patient's discharge abstract of each one of a set of ICD-9-CM codes was noted as a separate indicator variable. The set of ICD-9-CM codes for which indicators were defined for a given DRG was chosen as those present for at least 5 per cent of all patients in that DRG in the 1982 (index) data set; four- and five-digit ICD-9-CM codes were truncated to the first three digits. A patient's age, sex, and occurrence of death in hospital were also used as independent variables.

Clearly, the intercorrelations among laboratory tests and

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their transformations, as well as the large number of independent variables, could give rise to problems of multicollinearity, overfitting, and large computational demand. We chose the linear regression technique, partial least squares (PLS) regression, a latent variables procedure, because it is not much affected by any of these problems.[3]

For each DRG, PLS regression was used to define a (linear) predictor of LOS based on the data described above, using the 1982 data as the index set. To independently validate the LOS predictors thus derived, it was then applied to the 1981 data, which served as the validation set. The R^2 statistic was used as a measure of prediction performance. An R^2 of 0.00 was assigned to all DRGs which did not validate.

After the results using all 19 laboratory tests and all ICD-9-CM codes were obtained for the original data, several representative DRGs were examined to see if the number of laboratory tests in the regression could be reduced. Serum sodium, potassium, total carbon dioxide, and albumin appeared to contribute most of the information to the regressions examined. PLS was therefore re-run for all 48 DRGs in the index set and validated with the 1981 data set as above, using only those four laboratory tests, plus the demographic and ICD-9-CM data. Transformation, trimming, scaling and time-weighted mean were as before.

Results

The results of using the PLS regression defined on the clinical laboratory, demographic, and discharge data of the 1982 index set of patients to predict the LOS within each DRG for the 1981 validation set are summarized in Table 1. In the validation set, for 25 per cent of medical DRGs and 43 per cent of surgical DRGs, we were able to explain at least 10 per cent of the variability in length of stay by using additional clinical data. Overall, in 17 of 48 validation set DRGs examined, the ability of DRGs to predict length of stay was enhanced.

Table 2 summarizes the results when the four most "contributory" laboratory tests mentioned in the Methods section (sodium, potassium, bicarbonate, and albumin) were used selectively. In 23 of 48 validation set DRGs, the ability of DRGs to predict length of stay was improved by at least 10 per cent (50 per cent of medical DRGs and 46 per cent of surgical DRGs). These results were superior to those shown in Table 1, in which all 19 tests were used unselectively (see Discussion below).

Representative PLS regression coefficients created with the index set and used to predict LOS with the validation set for DRG 203, "malignancy of hepatobiliary system or pancreas," are shown in Table 3. Note, in particular, that higher sodium and albumin predict shorter LOS, while increasing age and surgical and radiological procedures predict longer LOS.

TABLE 1—LOS Prediction—19 Laboratory Tests Plus ICD-9-CM Codes and Demographic Data—Validation Set

	Number of DRGs (%)		
R^2 (%)	Medical	Surgical	All
>= 10	5 (25)	12 (43)	17 (35)
>= 20	2 (10)	9 (32)	11 (23)
>= 30	0 (0)	6 (21)	6 (13)
>= 40	0 (0)	1 (4)	1 (2)

TABLE 2—LOS Prediction—4 Laboratory Tests Plus ICD-9-CM Codes and Demographic Data—Validation Set

	Number of DRGs (%)		
R^2 (%)	Medical	Surgical	All
>= 10	10 (50)	13 (46)	23 (48)
>= 20	4 (20)	10 (36)	14 (29)
>= 30	3 (15)	4 (14)	7 (15)
>= 40	1 (5)	2 (7)	3 (6)

Discussion

The DRG system was created by a statistical partitioning process which created multiple new subgroups from "major diagnosis categories" if the reduction in length-of-stay variance (R^2) in the new groups was at least 2.5 per cent.[4] No severity-of-illness measure was used in the creation of these groups. As a result, it is not surprising that we, and other groups,[1,5,6] have shown that DRGs are heterogeneous with respect to hospital resource use.

Two widely discussed systems incorporating severity of illness into DRGs are Disease Staging [1] and the Severity of Illness Index.[5] The Disease Staging system initially required chart review to classify individual patients into stages. A computerized method based on ICD-9-CM codes was eventually developed, but the results of this system did not correlate well with the manual method.[7,8] Further, in one study,[1] the designers of the automated version of Disease Staging were able to explain a maximum of only 16 per cent of variance in total cost within "adjacent" DRGs (aggregated DRGs which ignore age, complications, and comorbidity) using their method.

Horn and colleagues report that their Severity of Illness Index allows major improvements in LOS prediction within DRGs.[9] Further, the system has been shown to be reproducible across institutions.[10] However, the Severity of Illness Index requires chart review for each patient by trained personnel and incorporates subjective ratings of illness severity. Because of its subjectivity, the system has been criticized as being "unusable" for prospective payment.[11] Additionally, the system is potentially "tautological", i.e., the rater could assume that a patient's length of stay is itself an indicator of that patient's severity of illness.[7] A computerized version of the Severity of Illness Index, i.e., Computerized Severity Index (CSI), has been developed[12] in response to these criticisms. This system, like our own, may potentially obviate the subjectivity criticisms. As with the computerized version of the Disease Staging system, however, Horn's CSI requires its own validation studies to document its equivalence to the manual method.

Our system is objective and data driven. The use of time-weighted mean laboratory results protects the system from unscrupulous manipulation; only the duration (relative to LOS) and degree of abnormality, not the frequency of testing nor LOS itself, enters into the predictions. Further, the data driven aspect of the system makes it adaptable both to changes over time and to differences among institutions, since the system itself derives the appropriate regression equations from the data.

TABLE 3—Regression Coefficients of Representative Explanatory Variables Used for Predicting LOS in DRG 203: Malignancy of hepatobiliary system or pancreas

Variables	Data Type	Coefficient	Sign	Explanation
Age	Demogr	0.18	+	Age of patient
NaMean	Clin Lab	0.29	−	Time-weighted-mean sodium
AlbMean	Clin Lab	0.25	−	Time-weighted-mean albumin
155	ICD-9 Dx	0.11	+	Mal. neopl. liver & intrahep. bile ducts
157	ICD-9 Dx	0.10	+	Mal. neopl. pancreas
529	ICD-9 Proc	0.14	+	Other operations on pancreas
922	ICD-9 Proc	0.20	+	Therapeutic radiol. and nuclear medicine

In performing our analyses, we chose to eliminate all patients who were LOS outliers, as defined by the criteria established by the Health Care Financing Administration (HCFA). Such patients are reimbursed differently under Medicare and have essentially been earmarked by HCFA as "distinct" within a DRG. By excluding LOS outliers, we have most likely penalized ourselves, since the LOS of non-outliers within a DRG would be more homogeneous than if they had been included, and, hence, there would be less variance to explain. Nonetheless, by deleting outliers, we have more stringently tested the ability of our system to improve the homogeneity of DRGs.

Logarithmic transformation of the laboratory test results tends to make the results more symmetric in distribution; data trimming moves extreme values towards the mean, eliminating the influence of "unusual" data elements. Although such techniques may be expected to help the predictor derived from fitting an index set be more robust, and hence perform well on an independent, future data set, the R^2 values derived directly from such a data set (here, the validation set) are the ultimate evidence of the validity of a predictor.

The four serum chemistries (sodium, potassium, total carbon dioxide, and albumin) found by our system to be the most important contributors to the regression model have precedent. The APACHE II system[13] found the serum sodium, potassium, and serum pH or bicarbonate at the time of admission to be predictive of resource use in intensive care patients. A study of nutrition-related variables in a group of surgical patients[14] showed that a low serum albumin was the most specific variable related to a length of stay greater than 10 days. Similarly, Baker and colleagues[15] found a high degree of correlation between serum albumin and clinical assessment of preoperative nutritional status in general surgical patients.

While it might seem contradictory that our results were better with only these four tests than with all 19 tests, it must be remembered that the results in question were from a separate validation set. There is nothing in the PLS procedure that guarantees that the final predictor chosen by it is, in fact, optimal among all possible predictors. Apparently, PLS exploited certain associations in the index set when defining its predictor for the 19-test case that simply failed to validate as well as expected.

We have used time-weighted average test results over a patient's entire hospital stay, rather than only admission laboratory results (as in APACHE II), because these values encompass a patient's overall metabolic derangements during his or her hospital stay. One criticism of incorporating tests beyond admission results is that this will not account for iatrogenically induced worsening of a patient's severity of illness. While such worsening would be expected to contribute to LOS, it might not be desirable to "reward" iatrogenic illness by additional compensation in a prospective payment plan. For this reason, we have begun to examine the use of time-weighted mean test results over shorter periods of time, such as the first 48 hours after admission.

A potential weakness of our system is that it depends in part on discharge abstract data, which are subject to coding errors[16] that were particularly problematic in the pre-DRG era. Since the advent of DRGs, however, with an expected decline in erroneous ICD-9-CM assignments,[16] this difficulty has lessened, and we are presently developing a more current data base to test our system further. In any event, by adding "noise," miscoding would decrease, not increase, the apparent value of discharge abstract data in predicting LOS. Another concern regarding the use of ICD-9-CM data is that the procedure codes may inappropriately enhance the predictive power of our system, if in fact they correlate with LOS as indicators of the process of care rather than as measures of severity of illness. We find, however, that more than 75 per cent of the explanatory power of our system comes from clinical laboratory data alone.

Another shortcoming of our methodology is that our results have been validated only with patient data at our hospital. A cross-hospital validation study would be desirable, and such a study is in progress.

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EPA Issues Interim Report on Pesticides in Groundwater

The US Environmental Protection Agency recently released an interim report, *Pesticides in Groundwater Data Base*, which summarizes available data on detections of pesticides in groundwater. Not to be confused with the National Pesticide Survey of Drinking Water Wells, the recently released report compiles data from approximately 150 groundwater monitoring studies carried out over the past 10 years by pesticide registrants, universities, and government agencies throughout the country. Findings of the interim report include:

- normal agricultural use has apparently led to residues of 46 different pesticides in the groundwater of 26 states;
- a total of 74 pesticides have been detected in the groundwater of 38 states from all sources, including normal agricultural use, accidental spills, residues from mixing, loading, and disposal of pesticides, back siphoning into wells during irrigation operations, etc;
- most of the pesticide concentrations detected in groundwater were at low levels, with median levels generally below EPA's health advisory levels for these pesticides.

While the interim report points to the potentially wide-ranging presence of pesticides at low levels in groundwater, the results do not offer a comprehensive picture of the severity of the problem of pesticides in drinking water on a nationwide basis. Many agricultural areas were not sampled by these studies and thus are not represented in the results. In addition, some of the studies included in the interim report were not limited to wells used for drinking water supplies but included shallow monitoring wells and wells used only for irrigation.

Copies of the interim report are available from the EPA Public Information Center, PM-211B, US EPA, 401 M Street SW, Washington, D.C. 20460. Telephone: (202) 475-7751.

The results of the National Pesticide Survey of Drinking Water Wells, scheduled for completion in 1990, are expected to meet the need for a statistically representative estimate of pesticide contamination of drinking water wells on a nationwide scale. The Survey will test 1,350 community and domestic wells for over 100 pesticides and breakdown products.

Physician Ratings of Appropriate Indications for Three Procedures: Theoretical Indications vs Indications Used in Practice

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Abstract: We previously reported substantial disagreement among expert physician panelists about the appropriateness of performing six medical and surgical procedures for a large number of theoretical indications. A recently completed community-based medical records study of about 4,500 patients who had one of three procedures—coronary angiography, upper gastrointestinal endoscopy, and carotid endarterectomy—shows that many of the theoretical indications are seldom or never used in practice. However, we find that there is also substantial disagreement (5, 25, or 32 per cent for angiography, endoscopy, or endarterectomy, respectively) about the appropriateness of indications used in actual cases if disagreement is defined by first discarding the two extreme of nine ratings, then looking for at least one rating near the bottom (1 to 3) and one near the top (7 to 9) of the 9-point scale. Patients should know that a substantial percentage of procedures are performed for indications about which expert physicians disagree. (*Am J Public Health* 1989; 79:445–447.)

Introduction

We previously reported on ratings by panels of physicians of the appropriateness of a large number of detailed indications for the use of six medical and surgical procedures.[1,2] The panelists assigned appropriateness ratings to indications in a process that was designed to encourage but not to require agreement. We found that, even in their final ratings, the panelists disagreed about the appropriateness of substantial fractions of the indications. One possible reason for the disagreement is that disagreed-upon indications may not occur in practice; our comprehensive lists included many indications that might be used only in theory.

Our purpose in this paper is to compare the panelists' agreement and disagreement about the theoretical indications with their agreement and disagreement about the indications actually assigned to cases in a recently completed medical records study.[3–5] That study determined the actual reasons for doing three of the procedures (coronary angiography, diagnostic upper gastrointestinal endoscopy, and carotid endarterectomy) in several thousand actual cases in five large areas of the United States. In this paper, we also compare the distribution by appropriateness or inappropriateness of the theoretical indications with that of the actual cases.

Methods

We have previously described in detail our two-stage modified Delphi process for obtaining appropriateness ratings.[1,2] The process used nationally representative panels consisting of nine expert physicians, including both generalists and specialists. The ratings on a nine-point scale ranged from extremely inappropriate (1), through equivocal (5), to extremely appropriate (9). Here we adopt the same four definitions of agreement that we previously used:

A9S: All nine of the ratings fell within a single three-point region—1 to 3, 4 to 6, or 7 to 9.

A9R: All nine of the ratings fell within *any* three-point range.

A7S: After discarding one extreme high and one extreme low rating, the remaining seven ratings all fell within a single three-point region—1 to 3, 4 to 6, or 7 to 9.

A7R: After discarding one extreme high and one extreme low rating, the remaining seven ratings all fell within *any* three-point range.

We also adopt the four previously used definitions of disagreement, and add a fifth. We used the fifth definition in our initial report of the medical records study:[4]

D9S: Considering all nine ratings, at least one was a 1 and at least one was a 9.

D9R: Considering all nine ratings, at least one fell in the lowest three point region (1 to 3) and at least one fell in the highest (7 to 9).

D7S: After discarding one extreme high and one extreme low rating, at least one of the remaining seven ratings was a 1 and at least one was a 9.

D7R: After discarding one extreme high and one extreme low rating, at least one of the remaining seven ratings fell in the lowest three point region (1 to 3) and at least one fell in the highest (7 to 9).

D5R: After discarding two extreme high and two extreme low ratings, at least one of the remaining five ratings fell in the lowest three point region (1 to 3) and at least one fell in the highest (7 to 9).

Deciding what to count as disagreement is necessarily subjective. We added definition D5R and used it in reporting the results of the medical records study[3] after examining patterns of individual panelist's ratings for representative indications. Many patterns that D7R would classify as disagreement, but D5R would not, seemed to us to show substantial clustering around the median rating—enough clustering so that we did not at that time want to count them as disagreement. We include all five definitions here so that readers can make their own choices.

In addition, we use our previous definitions of three appropriateness categories:[2]

Appropriate: a median rating of 7 to 9 without disagreement (using D7R to define disagreement for this purpose in this paper).

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Inappropriate: a median rating of 1 to 3 without disagreement.
Equivocal: a median rating of 4 to 6 without disagreement, or any median rating with disagreement.

During the medical records study, we abstracted hospital and office records for random samples averaging about 1,500 cases for each of the three procedures. The patient population consisted of people over age 65 covered by Medicare Part B. From the detailed abstracts, we determined which indication or indications applied to each case; if more than one indication applied, we chose the one with the highest median rating as the primary indication.

We compared the panelists' ratings of the full lists of indications that theoretically could occur, with their ratings of indications assigned to cases that actually did occur in practice. For each case, we noted the primary indication, and what category that indication fell into: agreement or disagreement; appropriate, equivocal, or inappropriate. We then calculated the per cent of cases whose primary indication fell into each category.

Note that the same ratings were used for the theoretical indications and for the actual cases; the panels did not separately rate the actual cases. The comparison we make in this paper is between an unweighted average for all of the indications on the theoretical list, and a weighted average in which each indication is weighted by the number of actual cases for that indication.

Results

Only a small number of indications for each procedure accounted for most of the cases found in the medical records study (see Appendix Table A). For example, the single most frequently occurring indication for angiography (preoperative evaluation of patients undergoing cardiac valvular surgery) accounted for 18 per cent of all cases. (See reference 3 for other frequently occurring indications.) The eight most frequently occurring indications (2.7 per cent of the 300 indications on the theoretical list) accounted for over half of the cases. The other two procedures exhibited similar concentration. We found only 20 to 40 per cent of the theoretical indications to occur at all in our samples of about 1,500 cases for each procedure.

Table 1 shows the percent of agreed-upon indications using each of the four definitions of agreement. The comparison of agreement on theoretical indications with agreement on indications assigned to actual cases is different for the three procedures.

- For angiography, the per cent of agreed-upon indications is about the same for the theoretical list as it is for the actual cases.
- For endoscopy, agreement on cases is substantially higher than is agreement on theoretical indications.
- For endarterectomy, agreement on cases is substantially *lower* than is agreement on theoretical indications.

As defined by A7R, agreement on the appropriateness of indications assigned to actual cases ranged from over 60 per cent for angiography to under 30 per cent for endarterectomy.

Table 2 shows disagreed-upon indications using each of the five definitions of disagreement. Disagreement on cases, compared with disagreement on the theoretical indications, was somewhat lower for angiography, about the same for endoscopy, and substantially higher for endarterectomy. As defined by D7R, disagreement on indications for actual cases was 5 per cent for angiography, but was 25 per cent for endoscopy and 32 per cent for endarterectomy.

Table 3 categorizes theoretical indications and indications assigned to actual cases by level of appropriateness. For all three procedures, the per cent of appropriate cases was higher than the per cent of appropriate theoretical indications—substantially higher for endoscopy and endarterectomy. The per cent of inappropriate cases was about the same

APPENDIX

TABLE A—Per Cent of Cases Done for the N Most Frequently Occurring Indications

Per Cent of Cases	Number and Per Cent of Indications: Angiography N	Angiography %	Endoscopy N	Endoscopy %	Endarterectomy N	Endarterectomy %
10	1	0.3	1	0.1	2	0.2
25	2	0.7	4	0.4	6	0.7
50	8	2.7	14	1.3	23	2.7
75	26	8.7	43	4.0	80	9.3
100	119	39.7	225	21.1	280	32.4

Note: N = number of indications
% = per cent of total theoretical indications:
300 for angiography,
1069 for endoscopy,
864 for endarterectomy

TABLE 1—Percent of Indications and Cases on Which Panelists Agreed Using Four Different Definitions of Agreement

Definition of Agreement		Coronary Angiography: Indic	Coronary Angiography: Cases	Endoscopy: Indic	Endoscopy: Cases	Carotid Endarterectomy: Indic	Carotid Endarterectomy: Cases
A9S	9 ratings, strict	28.0	29.5	25.4	39.5	40.9	16.4
	Standard error	(2.6)	(1.1)	(1.3)	(1.2)	(1.7)	(1.0)
A9R	9 ratings, relaxed	28.7	30.6	25.4	39.5	40.9	16.4
	Standard error	(2.6)	(1.1)	(1.3)	(1.2)	(1.7)	(1.0)
A7S	7 ratings, strict	50.0	55.5	41.3	53.4	53.4	27.3
	Standard error	(2.9)	(1.2)	(1.5)	(1.3)	(1.7)	(1.2)
A7R	7 ratings, relaxed	56.3	63.4	41.6	53.4	53.8	27.9
	Standard error	(2.9)	(1.2)	(1.5)	(1.3)	(1.7)	(1.2)

Notes: See text for full definitions of agreement.
'Indic' denotes the theoretical indications on the entire list rated by the panel. 'Cases' denotes actual cases to which indications were assigned in a field study of medical records.

TABLE 2—Percent of Indications on Which Panelists Disagreed Using Five Different Definitions of Disagreement

Definition of Disagreement		Coronary Angiography: Indic	Coronary Angiography: Cases	Endoscopy: Indic	Endoscopy: Cases	Carotid Endarterectomy: Indic	Carotid Endarterectomy: Cases
D9S	9 ratings, strict	2.0	0.1	30.2	26.6	14.9	21.7
	Standard error	(0.8)	(0.1)	(1.4)	(1.1)	(1.2)	(1.1)
D9R	9 ratings, relaxed	30.0	22.6	48.5	40.9	34.0	50.4
	Standard error	(2.7)	(1.0)	(1.5)	(1.2)	(1.6)	(1.4)
D7S	7 ratings, strict	0.3	0.0	7.0	6.3	2.3	2.7
	Standard error	(0.3)	(0.0)	(0.8)	(0.6)	(0.5)	(0.4)
D7R	7 ratings, relaxed	11.0	5.3	28.9	25.0	18.1	32.0
	Standard error	(1.8)	(0.5)	(1.4)	(1.1)	(1.3)	(1.3)
D5R	5 ratings, relaxed	1.3	0.9	11.4	5.2	7.0	6.6
	Standard error	(0.7)	(0.2)	(1.0)	(0.6)	(0.9)	(0.7)

Notes: See text for full definitions of disagreement.
'Indic' denotes the theoretical indications on the entire list related by the panel. 'Cases' denotes actual cases to which indications were assigned in a field study of medical records.

TABLE 3—Categorization of Indications by Appropriateness (%)

Appropriateness	Coronary Angiography		Endoscopy		Carotid Endarterectomy	
	Indic	Cases	Indic	Cases	Indic	Cases
Clearly inappropriate						
Median 1–3						
without disagreement	14.0	17.3	45.1	10.5	67.0	28.5
Standard error	(2.0)	(0.9)	(1.5)	(0.8)	(1.6)	(1.3)
Equivocal						
Median 4–6						
without disagreement	12.7	3.4	3.8	3.1	5.2	6.8
Any median						
with disagreement	11.0	5.3	29.2	25.0	18.1	32.0
Total	23.7	8.7	33.0	28.1	23.3	38.8
Standard error	(2.5)	(0.7)	(1.4)	(1.1)	(1.4)	(1.4)
Clearly appropriate						
Median 7–9						
without disagreement	62.3	74.0	21.9	61.3	9.7	32.7
Standard error	(2.8)	(1.1)	(1.3)	(1.2)	(1.0)	(1.3)

Notes. Disagreement means at least two 1–3 ratings and at least two 7–9 ratings. Indic denotes the theoretical indications on the entire list rated by the panel. Cases denotes actual cases to which indications were assigned in a field study of medical records.

as the per cent of inappropriate theoretical indications for angiography, and much lower for endoscopy and endarterectomy. The fraction of cases done for indications that we categorized as clearly inappropriate based on the panel ratings ranged from 10 per cent for endoscopy, through 17 per cent for angiography, up to nearly 30 per cent for endarterectomy.

The percentages in Table 3 differ somewhat from those that we have reported elsewhere.[3] The reason is that Table 3 is based on disagreement defined by D7R, whereas the earlier report used D5R. We adopt the more easily satisfied definition here in order to provide a new perspective on the appropriateness results previously reported.

Discussion

Disagreement among expert physicians about appropriate indications for doing angiography, endoscopy, and endarterectomy is not limited to theoretical indications that are never done in practice. We found that there was generally almost as much or even more disagreement about the appropriateness of indications assigned to actual cases than there was about the complete lists of theoretical indications. Disagreement about either indications or cases was minimal for coronary angiography on most of our definitions. In contrast, there was substantial disagreement about both the theoretical indications and the indications actually used for endoscopy and carotid endarterectomy. For example, 27 per cent of endoscopies were done for reasons that at least one out of nine panelists rated 1 (extremely inappropriate) and at least one rated 9 (extremely appropriate), and thus count as disagreement on definition D9S. Also, 32 per cent of endarterectomies were done for indications that at least two panelists rated 3 or lower and at least two panelists rated 7 or higher (disagreement on definition D7R).

There is generally more difference between the theoretical indications and the indications assigned to actual cases when they are categorized by appropriateness (Table 3), than when they are categorized by agreement or disagreement (Tables 1 and 2). Particularly for endoscopy and endarterectomy, inappropriate indications occur frequently on the theoretical lists, but most of them are not used in practice.

The high percentages of disagreement about the appropriateness of doing endoscopies and endarterectomies for indications that were assigned to actual cases reflect the real uncertainty about the efficacy of these procedures in many situations.[6,7] Clinical trials may help to lessen the uncertainty in the future, but can never eliminate it entirely. Physicians should carefully weigh the advisability of doing a procedure for indications that lack firm scientific justification. If they do believe that the procedure is advisable in such circumstances, they should make sure that their patients understand that other physicians might disagree. Also, patients should attempt to ascertain and to understand the reasons for their doctor's recommendations before submitting themselves to expensive, invasive procedures.

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Beliefs about AIDS as Determinants of Preventive Practices and of Support for Coercive Measures

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Abstract: We performed a telephone survey to explore relations between knowledge, beliefs (as defined in the Health Belief Model) and reported AIDS-preventive practices in a sample of 1,072 persons ages 18-65, living in the Montreal health region.

AIDS-preventive practices were more frequent among the young or single, and among those with one of four health beliefs: perceiving oneself as particularly susceptible to AIDS, perceiving the disease as particularly severe, perceiving it as particularly amenable to prevention, and having a strong general health motivation. Support for coercive measures to control the AIDS epidemic was widespread but was stronger among the less educated, married people, and those with a high level of one of the following beliefs about AIDS: perceived severity, susceptibility, curability, or barriers to treatment.

AIDS-preventive practices and support for coercion under epidemic conditions share their most important modifiable determinants: perceived severity of AIDS and perceived susceptibility to it. This finding suggests that emphasizing them, as is done so often in public educational campaigns about AIDS, may indeed promote preventive practices, but may also unwittingly increase support for coercive measures toward people with the disease or at high risk of it. (*Am J Public Health* 1989; 79:448–452.)

Introduction

Although the incidence of the acquired immunodeficiency syndrome (AIDS) has remained low outside of the recognized high-risk groups,[1] concern about an eventual increase in transmission among groups now at low risk has led to large-scale information programs promoting AIDS-preventive practices in the general population. Studies of practices intended to reduce the risk of AIDS in various groups have been extensively reviewed.[2] Although information is available about these in some high-risk groups (homosexual men, intravenous drug users, hemophiliacs, single heterosexual men, heterosexual partners of HIV (human immunodeficiency virus)-positive persons, adolescents and young adults, prostitutes and their clients) relatively little is known about preventive practices in the general population.

The theoretical model that has guided the design of any published studies,[3] the Health Belief Model (HBM),[4] makes it indeed attractive as a theoretical approach to AIDS-related behavior. Scales of demonstrated validity and reliability have been developed, using standard psychometric methods, for measuring health beliefs.[5–7]

The objectives of the present study were to:

- describe how knowledge, beliefs and practices related to AIDS are distributed among the general population of the Montreal health region, including all French- and English-speaking persons 18-65 years of age;
- identify associations between AIDS-preventive practices and sociodemographic characteristics, knowledge and health beliefs of this population.

Methods

The Questionnaire*

The sociodemographic variables included sex, age, educational level, mother tongue, marital status, and total household income.

Knowledge was measured by scales on AIDS symptoms (six items), modes of transmission (six items), and high-risk groups (eight items). Each scale had equal numbers of true and false statements; agreeing with a true statement added one point to the score for the scale, as did disagreeing with a false one; the opposite choices resulted in the loss of a point.

Health beliefs were measured by six Likert-type scales: perceived severity of AIDS (five statements with which the respondent had to express his degree of agreement or disagreement), perceived susceptibility to it (five statements), perceived benefits of its prevention (six statements), perceived benefits of its treatment (five statements), perceived barriers to its treatment (two statements), and general health motivation (six statements). The scales measuring severity, susceptibility, and general health motivation were borrowed from Champion,[7] substituting "AIDS" for "breast cancer"; the others were created by the investigator on the same model. All scales showed a high degree of internal consistency (in a pretest using a convenience sample of persons attending a travelers' immunization clinic), after statements poorly correlated with the total score for their scale had been eliminated.

AIDS-preventive practices were probed by 14 questions: five on changes in sexual behavior, four on health care sought or avoided, and one each on substance abuse, avoidance of social contacts, precautions at work, condom use, and avoidance of blood donations.

Support for coercive measures was assessed by questions on whether the respondent would support each of four types of coercive actions by the government to control the AIDS epidemic.

The Survey

The data collection was carried out by a major telephone polling agency using random digit dialing. Standard procedures determined the generation of telephone numbers, the random selection of the respondent within the household (no substitution allowed), the number of calls to be made in case of non-response, and call-backs for quality control.

The univariate statistical analyses were done by contingency tables, after all variables had been dichotomized. The 95% confidence intervals (CI) of the odds ratios (OR) were calculated by the test-based method.[8] All potential determinants of preventive practices and of support for coercive measures were also subjected (separately) to logistic regres-

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*Details available on request to author.

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sion analyses, using program BMDPLR with the default options in force. An attempt was made to introduce into the equation interactions between the most important terms. ORs and their 95% CIs were calculated from the regression coefficients in the usual way.[8] The comparison between two ORs when one is larger and the other smaller than unity is made easier by taking the reciprocal of one of them (thus inverting the high- and low-risk categories).

Results

Response Rate

A total of 2,677 telephone numbers were generated, of which 64 per cent (1,712) belonged to the target population (i.e., the numbers were in service, residential, and reached a household with at least one member meeting the age and language criteria). Of the 1,712 eligible respondents, 63 per cent (1,072) were interviewed. Non-response was due to refusal or repeated absence of the respondent selected within the household (29 per cent and 4 per cent, respectively) or to the impossibility of reaching anyone at the number (4 per cent). The interviews were done between March 25 and May 24, 1987.

Description of the Sample

Forty-two per cent of respondents were men and 58 per cent women; 18 per cent were between ages 18 and 24, 16 per cent between 25 and 29, 25 per cent between 30 and 39, 18 per cent between 40 and 49, and 23 per cent were age 50 or over. The mother tongue was French for 70 per cent, English for 18 per cent, and other for 12 per cent; 79 per cent of interviews were carried out in French, the rest in English. Fifty-two per cent of the respondents were married or living in a common-law relationship, 32 per cent were single, 13 per cent were separated or divorced, and 3 per cent were widowed. The highest education level attained was grade school for 7 per cent, high school for 43 per cent, junior college for 20 per cent, and university for 30 per cent. Total household income was below $20,000 (Canadian) for 31 per cent, and above $40,000 for 26 per cent.

Knowledge

Almost all respondents (98.6 per cent) had heard of AIDS. Knowledge of the modes of transmission was highest, with 81 per cent of respondents having the maximum possible score; knowledge of high-risk groups was intermediate and that of AIDS symptoms lowest, its mean being in the middle of the possible range. All measures of knowledge about AIDS were higher among the more educated, as has been reported for other aspects of health knowledge.[9]

Health Beliefs

As a whole, the distributions of health beliefs were not as skewed as those of knowledge. Only four strong associations between health beliefs and sociodemographic characteristics were detected: women have a higher general health motivation than men, unmarried persons (even when controlling for younger age) have a higher perceived susceptibility to AIDS than married persons, better-educated and French-speaking persons have a higher perceived preventability of AIDS. Apart from these instances, beliefs about AIDS appear to be uniformly distributed among basic social groups.

AIDS-Preventive Practices

Sixty-nine per cent of respondents denied all of the 14 practices asked about, 15 per cent acknowledged one, 8 per cent, two, 4 per cent three, and 4 per cent between four and nine; in all, 31 per cent acknowledged at least one practice. The frequencies of the most common practices were:

- voluntarily decreasing the number of sex partners (13 per cent);
- changing one's sex behavior (12 per cent);
- using a condom to avoid the transmission of AIDS (10 per cent);
- avoiding social contact with certain people or groups, mostly homosexuals, prostitutes, and casual sex partners (6 per cent);
- taking precautions at work, such as health care workers wearing gloves, masks or blouses, washing hands, handling needles carefully, taking precautions in relation to eating or avoiding certain people (6 per cent);
- avoiding certain types of sexual behavior, including vaginal intercourse (1.7 per cent), oral intercourse (2.2 per cent), and anal intercourse (2.4 per cent);
- avoiding giving blood (4 per cent);
- visiting a doctor because of AIDS (2 per cent);
- refusing to receive a blood transfusion (1 per cent)
- five other practices (less than 1 per cent each).

Unfortunately, too little is known about the respondents to assess whether each practice acknowledged by each respondent was indicated or superfluous for this individual, except for the completely ineffective practices.

Support for Coercion

Four questions on the issue of coercion and the answers given were as follows:

"Do you think children with AIDS should attend school in the same way as other children?" Yes 70%, no 20%, don't know 10%.

"Do you think the government should declare illegal certain kinds of sexual activity, if this helps fight the AIDS epidemic?" Yes 40%, no 50%, don't know 10%. Types of sexual activity most often mentioned: homosexuality (14% of all 1072 respondents) and prostitution (7%).

"Do you think that people with AIDS should be prevented from holding certain kinds of jobs, if this helps fight the AIDS epidemic?" Yes 44%, no 45%, don't know 10%. Jobs most often mentioned: those in hospitals or in the medical field (13%), involving contact with the public (6%), in the food industry or in restaurants (6%).

"Do you think that certain groups of people should be forced to take a test to find out if they have AIDS, if this helps fight the AIDS epidemic?" Yes 74%, no 21%, don't know 4%. In all, 23 groups were named; most commonly: homosexuals or bisexuals (23% of all respondents), everyone (11%), and prostitutes (5%).

The number of questions for which the respondent chose the "coercive" answer can be considered a variable indicative of the strength of his support for coercive measures to control AIDS, on a scale of 0 to 4. The percentages of respondents in each category were respectively 14, 31, 27, 21 and 8 per cent. In subsequent analyses, subjects in the last two categories will be considered to support coercion, those in the first two not to support it, and those in the third will be left out of the analyses.

Determinants of Preventive Practices and of Support for Coercion

Table 1 presents the percentage of respondents having carried out at least one preventive practice, when each potential determinant is present and absent, with the corresponding ORs and their 95 per cent CIs. Being young, male, single, of high education, of low income, perceiving AIDS as

TABLE 1—The Determinants of AIDS-Preventive Behavior*

Determinants	Present	Absent	OR	(95% CI)
Demographics				
Age (< 35)	38.4	23.2	2.06	(1.58, 2.70)
Sex (male)	35.5	27.1	1.48	(1.13, 1.92)
Language (French)	29.1	34.2	0.79	(0.60, 1.05)
Education (college and up)	35.2	26.0	1.55	(1.19, 2.01)
Marital status (unmarried)	45.0	17.2	3.93	(2.99, 5.17)
Income (< $30,000)	36.0	25.7	1.63	(1.24, 2.14)
Knowledge (above median)				
Symptoms	31.6	29.4	1.11	(0.85, 1.45)
Transmission	31.2	28.4	1.14	(0.81, 1.60)
High-risk groups	32.8	28.3	1.24	(0.95, 1.61)
Health belief scores (above median)				
Severity	37.3	24.3	1.85	(1.41, 2.40)
Susceptibility	41.1	21.0	2.62	(2.00, 3.43)
Preventability	33.5	28.4	1.27	(0.97, 1.65)
Curability	30.6	30.7	1.00	(0.76, 1.32)
Barriers	29.1	32.6	0.85	(0.65, 1.10)
General Health Motivation	33.5	27.2	1.35	(1.03, 1.75)

*Percentage of subjects having carried out at least one preventive practice, in the presence and absence of each potential determinant, with the corresponding odds ratio and 95% confidence interval. (N = 1050).

a particularly severe illness, perceiving oneself as particularly susceptible to AIDS, or having a high general health motivation are all significantly associated with having carried out at least one preventive practice. Interestingly, knowledge is not a determinant.

Table 2 presents in a similar way the determinants of support for coercion. Support is more common in those who are older, married, of low education or low income, with a poor knowledge of the modes of transmission or of high-risk groups or having a high level of one or more of these beliefs about AIDS: perceived severity, susceptibility, preventability, and barriers to treatment.

There also are many associations among the determinants themselves, so that the effects presented in the two previous tables cannot be considered independent of each other. Logistic regression was used to unravel these relations. The resulting models are presented in Table 3.

The determinants of having carried out a preventive practice no longer include education or income, probably because these are strongly associated with age. Sex is also eliminated, by a hair's breadth, probably because of its association with the health beliefs. Having a high perceived preventability of AIDS becomes significant.

The determinants of support for coercion no longer include age, income, or any of the knowledge measures, probably because of their association with education. Perceived preventability loses its significance, while perceived curability becomes significant.

In neither case was there any interaction between the most important terms of the regression equation.

One notes that the two outcomes share their most important modifiable determinants (perceived severity and susceptibility). Indeed, the two outcomes are associated: a support for coercion is present among 45 per cent of respondents who have carried out one or more practice and in 37 per cent of those who have not done so.

TABLE 2—The Determinants of Support for Coercive Measures*

Determinants	Present	Absent	OR	(95% CI)
Demographics				
Age (< 35)	36.1	43.9	0.72	(0.54, 0.97)
Sex (male)	39.6	40.0	0.98	(0.73, 1.31)
Language (French)	41.5	36.1	1.26	(0.91, 1.72)
Education (college and up)	28.3	51.8	0.37	(0.27, 0.49)
Marital status (unmarried)	35.5	44.0	0.70	(0.52, 0.93)
Income (< $30,000)	45.0	33.2	1.64	(1.21, 2.22)
Knowledge (above median)				
Symptoms	37.0	43.8	0.76	(0.56, 1.01)
Transmission	36.4	54.4	0.48	(0.34, 0.69)
High-risk groups	34.5	46.0	0.62	(0.46, 0.83)
Health belief scores (above median)				
Severity	52.8	27.2	2.98	(2.22, 4.01)
Susceptibility	48.8	31.6	2.06	(1.54, 2.76)
Preventability	35.3	43.6	0.70	(0.52, 0.94)
Curability	42.9	36.7	1.29	(0.97, 1.73)
Barriers	44.8	33.5	1.61	(1.20, 2.16)
General Health Motivation	41.8	37.7	1.19	(0.89, 1.59)

*Percentage of subjects with a high degree of support, in the presence and absence of each potential predictor, with the corresponding odds ratio and 95% confidence interval. (N = 723)

One notes that the two outcomes share their most important modifiable determinants (perceived severity and susceptibility). Indeed, the two outcomes are associated: a support for coercion is present among 45 per cent of respondents who have carried out one or more practice and in 37 per cent of those who have not done so.

Discussion

A response rate of 63 per cent is close to the average rate of 65 per cent reached by the agency for telephone surveys in the Montreal region; how non-respondents differ from respondents, and whether this could have affected the results obtained, is unknown. Most respondents were French-speaking, but language influenced neither the frequency of practices nor support for coercion, so that the results appear generalizable to English-speaking populations.

A purely cross-sectional study cannot establish the time-relations in the associations it detects, and thus provides no grounds for deciding whether the associations are causal or not. Neither knowledge nor beliefs necessarily precede the behavior with which they are associated. They *may* have helped bring it about; or they may have resulted from it, people selectively acquiring knowledge and beliefs consistent with behavior that has already been determined by quite different influences. The likelihood of this having happened is

TABLE 3—Determinants of AIDS-Preventive Practices and of Support for Coercive Measures, as Identified by Logistic Regression. Adjusted Odds Ratios and 95% Confidence Intervals

Determinants	Preventive Practices Adjuted OR (95% CI)	Support for Coercion Adjusted OR (95% CI)
Demographics		
Age (<35)	1.46 (1.07, 2.00)	0.85 (0.59, 1.21)
Sex (male)	1.35 (1.00, 1.84)	1.08 (0.77, 1.52)
Language (French)	0.78 (0.56, 1.09)	1.21 (0.83, 1.77)
Education (college and up)	1.33 (0.96, 1.83)	0.45 (0.31, 0.63)
Marital status (unmarried)	3.20 (2.33, 4.40)	0.65 (0.46, 0.93)
Income (<$30,000)	1.33 (0.97, 1.83)	1.41 (0.99, 2.02)
Knowledge (above median)		
Symptoms	1.10 (0.81, 1.50)	0.95 (0.68, 1.34)
Transmission	0.96 (0.64, 1.45)	0.74 (0.48, 1.14)
High-risk groups	1.29 (0.94, 1.76)	0.79 (0.56, 1.11)
Health belief scores (above median)		
Severity	1.68 (1.24, 2.28)	3.12 (2.22, 4.39)
Susceptibility	2.38 (1.74, 3.26)	1.53 (1.08, 2.16)
Preventability	1.38 (1.01, 1.89)	1.06 (0.75, 1.50)
Curability	0.93 (0.69, 1.26)	1.45 (1.04, 2.04)
Barriers	0.82 (0.60, 1.11)	1.60 (1.14, 2.25)
General Health Motivation	1.54 (1.13, 2.10)	1.02 (0.73, 1.44)

perhaps less for a new disease like AIDS than for one people have known for most of their lives.

It may seem surprising that 31 per cent of respondents reported some behavioral adaptation to the AIDS epidemic, but the proportion is in fact close to that found in studies of adolescents[10] or of the Canadian population.[11]

Unfounded practices are less frequent than appropriate ones, although not as rare as one would wish. The most frequent ones are avoidance of social contacts (6 per cent) and of donating blood (4 per cent). These remain pertinent targets for specific educational efforts.

The formulation of the four questions on support for coercion calls for some comments. The question on school attendance is the only one for which a "yes" answer (which, everything else being equal, respondents prefer to give—the "acquiescence response set") meant a non-coercive stand. This may be a reason why the answers to it favored coercion much less than those to the other questions. A more encouraging reason might be that the issue of school attendance has arisen concretely, has been discussed in the media, and has sometimes given rise to constructive and compassionate reactions. The other three questions all end with "if this helps fight the AIDS epidemic." This clause (including the emotionally charged word "epidemic" which would usually be avoided in questionnaire formulation) was added in order to simulate these measures being advocated. Presumably, they would be touted as necessary to fight the epidemic, not as a novel way to oppress people. Hence the answers should be interpreted as a measure of latent rather than actual support for coercion, except perhaps for respondents who believe themselves to be in the midst of an AIDS epidemic.

The high level of support for coercive measures detected in this survey corresponds to the findings of previous studies. Mandatory testing is widely supported, whether it be for everyone (52 per cent of Americans in favor)[12] or for high-risk groups (about two-thirds of Canadians in favor).[11] However, in this study, there is little agreement on who these measures should apply to. Such "free floating" support may be more easily counteracted if it is not given time to crystallize around specific measures against specific groups.

A methodological note of caution—odds ratios are used as measures of effect because they can be computed from logistic regression coefficients. However, neither dependent variable is rare (practices in 31 per cent of 1,050 respondents, support for coercion in 40 per cent of 723 subjects in this analysis), so that the OR tends to overestimate the size of the effect. The magnitude of this phenomenon can be estimated by taking the ratios of the percentages presented in Tables 1 and 2 and comparing these rate ratios with the corresponding ORs.

The absence of significant interactions between determinants indicates that the effect of a given change in health beliefs would not be significantly different in any subgroup (whether defined by age, sex, marital status, or income). Thus this survey provides no evidence suggesting that one should try and modify some beliefs in some groups and other beliefs in other groups. Obviously, there will be groups with strong sets of beliefs about AIDS which differ from those of the general population (such as gay men, hemophiliacs or IV drug users) but the survey was not aimed at them.

As for the groups that the study did identify, it is easy to understand why the young and the unmarried have adopted AIDS-preventive practices more frequently than the others (even though information on their usual sexual activities is unavailable); this corresponds to the findings of American surveys.[13] The higher frequency in men (32 per cent) compared to women (24 per cent) is puzzling, other determinants being equal as estimated from the regression equation (unadjusted figures are 36 and 27 per cent). Assuming that 5 per cent of men are homosexual, that 100 per cent of homosexual men have adopted at least one preventive practice and that the same proportion of heterosexual men and women have adopted a preventive practice, one can compute that in all 28 per cent of men would have adopted a practice:

$$(1.00 \times 0.05) + (0.24 \times 0.95) = 0.28.$$

This is the mid-point between the observed proportions in men and women. Thus homosexuality in men can explain up to half of the difference between the sexes, but the other half is unaccounted for, since other measured determinants have been taken into account in calculating these percentages. Thus there appears to be a lag in the acceptance of preventive practices by women in this survey, contrary to what is reported in the US.[13]

It has been suggested that willingness to support coercive measures is associated with having false information about the transmission of HIV.[14] In this survey, having agreed with false statements about transmission is indeed strongly predictive of support for coercive measures, in univariate analyses, but, as in the case of knowledge scores, the effect does not hold up in multivariate analyses. These results fail to confirm the existence of a relation between misconceptions about transmission and support for coercion.

Finally, the finding that preventive practices and support for coercion share the same important modifiable determinants is probably the most thought-provoking result of this study. Many health education messages aim precisely at creating fear of AIDS and feelings of vulnerability to it. Such messages may increase support for coercion as well as foster preventive behavior. Let us not forget the warnings about the counterproductive effects of excessive levels of fear,[15] and about the importance of complementing messages on the severity and probability of an untoward health event with information about the effectiveness of preventive action[16] and with positive reinforcement of desirable forms of behavior.[17] In this respect, it is worth noting that a strong belief in the preventability of AIDS was associated (although weakly) with having adopted at least one AIDS-preventive practice, but not with supporting coercive measures.

Clearly, the many health educational efforts made to contain the AIDS epidemic need to be guided by a better knowledge of all their consequences, whether intended or not. The Health Belief Model appears useful for generating this knowledge.

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Accuracy of Cholesterol Level Tests Should be Double-Checked

A single test may not accurately reflect a person's cholesterol level, and errors can result in significant differences in medical treatment, according to the February issue of the *Tufts University Diet & Nutrition Letter*. Part of the problem is that many laboratories as well as manufacturers of cholesterol testing equipment have not been using the same test methods, and/or have not been setting their equipment to the standard established by the National Cholesterol Education Program of the National Institutes of Health.

A 1985 study done by the College of American Pathologists found that readings may be off as much as 15 percent, meaning that someone with a true reading of 220 milligrams per deciliter of blood could be told their cholesterol level is from 187 to 253. With a level of 200 considered desirable, misreadings can mean the difference between a clean bill of health and a diagnosis of being at risk for heart disease. But laboratories and manufacturers are making changes. And by asking the right questions, individuals can ensure the accuracy of their own readings before deciding on new dietary regimes.

Herbert K. Naito, chairman of the Laboratory Standardization Panel of the National Cholesterol Education Program, noted that the dozen manufacturers who produce about 75 percent of cholesterol testing equipment in the United States are calibrating their equipment according to national standards.

In addition, the Centers for Disease Control in Atlanta have set up a network of reference laboratories that use the accepted standardization so that other labs can check the accuracy of their analyses.

Preventing AIDS and Other STDs through Condom Promotion: A Patient Education Intervention

MILDRED Z. SOLOMON, MA, AND WILLIAM DEJONG, PHD

Abstract: We report on two studies that assessed the impact of a soap-opera style videotape on inner-city STD (sexually transmitted disease) patients' knowledge about and attitudes toward condom use, and willingness to redeem coupons for free condoms. Subjects in the first study who viewed the videotape (and participated in a brief oral recall session) had higher knowledge scores and more accepting attitudes than subjects who did not (knowledge test means of 11.1 versus 7.9, attitude index means of 13.0 versus 11.3). The intervention was most effective among those who were relatively poorly educated and, to a lesser extent, among those who reported less frequent use of condoms and fewer sex partners. In the second study, intervention group subjects were more likely than control group subjects to redeem coupons. Both groups exhibited a high level of interest in the free condoms. We argue that education *and* accessibility to free condoms can increase condom use and that health care providers have a vital role in promoting this form of STD prevention. (*Am J Public Health* 1989; 79:453–458.)

Introduction

Condom use appears to be one of the most important weapons in the public health fight against AIDS (acquired immunodeficiency syndrome).[1-3] Campaigns to promote condom use have been launched within the gay community[4] and among prostitutes,[5] but much more needs to be done, especially to reach the drug-using population, female partners of drug-users and bisexuals, individuals with multiple sex partners, and minority populations.[2-5]

Although the AIDS crisis has spotlighted national attention on the need for condoms, the US Public Health Service called for an increase in condom use in 1980,[6] well before AIDS was the national public health priority it is now. Both laboratory studies of condom permeability and epidemiological field studies of disease incidence among those who do and do not use condoms have documented that condoms are effective against a wide variety of sexually transmitted diseases (STDs) including gonorrhea,[7,8] nongonococcal urethritis,[9,10] trichomoniasis,[10,11] and herpes genitalis,[12,13] as well as AIDS.[14,15] Although one recent review of the literature suggests that clinical evidence for the condom's effectiveness against AIDS is less than conclusive,[16] public health authorities have called for the use of condoms by individuals at high risk.[17] Despite their promotion, however, condoms are underutilized, especially among individuals who are most at risk, such as persons with multiple sex partners,[10,18] or those who have been previously treated for an STD.[19]

Historically, condom use in the US has been associated with extramarital sex, prostitution, and promiscuity.[20-22] In addition, condoms are perceived by both men and women to impair the pleasure of sexual intercourse.[23-25] Furthermore, introducing condom use can entail a difficult interpersonal negotiation that can result in distrust and embarrassment.[26]

Prior to the AIDS crisis, there were very few concerted efforts by the public health community to promote condom use in the United States. The paucity of domestic condom promotion efforts stands in sharp contrast to the large-scale public education campaigns undertaken abroad in both developing and industrialized countries[20,27-29] for birth control purposes and for STD prevention.[30] These campaigns have often employed social marketing techniques, promoting positive images of condoms while providing greater accessibility to them through, for example, more convenient purchase outlets and subsidized prices. Recent pressure on the national television networks in the US to air condom advertisements suggests that, with continuing public concern about AIDS, we may eventually see large-scale social marketing of condoms in this country.

We conducted two studies to ascertain the effects of a patient education videotape and an oral recall session on patient's knowledge about the need for condoms and their proper use, attitudes toward condoms, and motivation to redeem coupons for free condoms. We also studied whether the intervention was more successful with different patient subgroups. The procedures for both studies were reviewed and approved by the Institutional Review Board of Education Development Center, Inc.

The Videotape

The videotape, entitled "Let's Do Something Different," follows the story of Diane, a young Black woman who has had two cases of gonorrhea and recently returned from the hospital after a bout of pelvic inflammatory disease. Through the support of her brother and a girlfriend, Diane comes to realize that she can persuade her boyfriend to use condoms, and we see her communicate with him successfully, with good grace and humor. Several features of the videotape merit attention:

- *The videotape portrays using condoms as socially acceptable, normative behavior.* The protagonist's growing confidence that she can influence her lover is due to the persistent advice of her brother and girlfriend, both of whom are sexually attractive and sexually active, yet committed to the values of health and sexual responsibility. The brother, and later the converted boyfriend, serve as role models for male viewers by providing examples of men who understand the severity of STDs and who are comfortable expressing their love and concern for a woman they care about.
- *The focus is on the development of interpersonal and communication skills.* The videotape presents only the most essential factual information about disease etiology and transmission. Furthermore, "safe sex" is promoted not by graphic depictions of sex acts, but by modeling effective communication styles and conflict

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resolution—those interpersonal skills required for negotiating safe sex with one's partners.

- *The protagonist is a woman.* A consequence of asymptomatic venereal infection in women, pelvic inflammatory disease (which often results in sterility) is endemic in the U.S., with 857,000 cases reported per year.[30] Black women bear a disproportionate burden, with a rate nearly three times that for White women.[31] We chose a female protagonist and a feminine point of view for our narrative in order to highlight the heavy toll that unprotected intercourse takes on the reproductive health of women. We hoped that, by coming to understand the consequences of unprotected intercourse on women and children, men (who constitute the vast majority of patients seen in most public clinics) would be more motivated to change their own attitudes and behaviors.
- *It strives to create a sense of identification within viewers.* Since the videotape is intended primarily for use in public clinics that serve a predominantly Black population, most of the characters are Black and they express many of the same attitudes, beliefs, and values that our background research revealed to be important to this patient population. The script purposely avoids the use of medical terms, using instead the argot of the target group.
- *The videotape attempts to make condoms more sexually appealing.* Although it acknowledges their disadvantages, the videotape refers to the ways in which condoms can enhance sexual pleasure—by prolonging intercourse[32]; by their slight tourniquet effect on the superficial veins of the penis, which may help sustain erection[32]; and by giving both partners peace of mind.[23] Also, because of the widespread complaint that condoms are disruptive, the videotape suggests that putting the condom on can be integrated into foreplay. "Eroticizing" the condom in this way is consistent with the tenets of social marketing,[20,33,34] which hold that target behaviors should be associated with existing values and desires already held by the audience.

STUDY 1: KNOWLEDGE AND ATTITUDES

Method

During the spring of 1986, clinicians at the STD Clinic at Boston City Hospital referred to the study all patients seen at the clinic except those who were under 18 years of age, were non-English speaking, had returned for a test-of-cure examination, or needed to be examined by a specialist. A total of 214 patients were referred. Of those, our research assistant had to pass over 59 to avoid excessive clinic backlog; eight were absent when called; four were non-English speaking; 10 were under 18 years of age and could not be in the study; and four had been seen previously.

While each potential subject waited to receive his or her laboratory test results, the research assistant explained the purpose of the study and solicited his or her informed consent to participate. Of the 129 patients asked, 103 (79.8 per cent) agreed to participate. Half of the consenting subjects were assigned randomly to see the videotape (intervention group, n=51); half were not (control group, n=51). All subjects in the intervention group viewed the videotape individually.

Measurement

All subjects were asked several questions about their demographic background, STD history, number of sexual partners in the past month, and estimated frequency of condom use (never, almost never, sometimes, most times, all times).

Intervention group subjects were shown the videotape and, immediately after viewing, were asked to recall its main points. As the subjects talked, the research assistant checked off on a master list which points were recalled and wrote out additional points mentioned if they did not appear on the checklist; general probes (e.g., Anything else? Can you think of anything more?) were used until subjects had nothing more to add.

Both intervention and control group subjects took a 13-item true-false test, administered by the research assistant. All of the test items were derived from the content of the videotape. Item number 13 had a negative item-total correlation and was omitted from further analysis. The remaining 12 items showed good internal reliability (coefficient alpha =.64). Subjects were also asked 14 additional true-false questions designed to measure their attitudes toward condom use. An attitude index derived from these 14 true-false items showed good internal reliability (coefficient alpha =.66).

Subjects then were asked the following open-ended question: "If a person wants to convince his/her sex partners to use condoms, what are some things that person could do or say?"

Patient Profile

Study subjects were predominantly male (81.6 per cent), single (74.8 per cent), Black (79.6 per cent), and native-born (78.4 per cent). Subjects ranged in age between 18 and 51 years (median = 24 years); 77.5 per cent were 30 years or younger; 24.3 per cent had completed only junior high, 36.9 per cent had completed only high school, and 36.9 per cent had attended college or a vocational/technical school.

Sixty-four per cent of the subjects had a prior history of sexually transmitted disease, and about one-fifth reported a diagnosis within the last six months. Just under half of the subjects reported having more than one sex partner during the past month. Fifty-eight per cent of subjects reported that they had never or almost never used condoms prior to their present visit.

Intervention and control groups did not differ with respect to sex, marital status, race, age, education, birthplace, prior STD history, number of sex partners in the past month, or previous condom use.

Results

Patients' Knowledge and Beliefs

Table 1 presents the percentage of intervention and control group subjects who responded correctly for each item of the true-false test. Subjects who saw the videotape scored higher than subjects who did not: 70.6 per cent of the intervention group correctly answered 11 or all 12 items; only 7.7 per cent of the control group subjects scored that well.

Patients' Attitudes

Table 2 presents the percentage of intervention and control group patients who indicated a favorable response for each item of the attitude index. The mean score on the attitude index was higher for the intervention group (13.0) than the control group (11.3), indicating a more favorable attitude toward condom use among subjects who saw the videotape. While 70.6 per cent of intervention group subjects answered 13 or 14 of the attitude items favorably, only 32.7 per cent of the control group patients did so.

Strategies to Persuade Partners to Use Condoms

As shown in Table 3, intervention group subjects cited more strategies for persuading partners to use condoms, exclusive of citing the condom's effectiveness in preventing pregnancy and disease. The most commonly mentioned alternative strategies included: refusing to have sex; and citing that using condoms could be fun and did not have to interfere with lovemaking.

TABLE 1—Percentage of Patients with Correct Response on Each Item of the Knowledge Test by Experimental Condition

Knowledge Test Items*	% Correct Responses			
	Intervention Group	Control Group	Difference	(95% CI of Difference)
1. Using a condom can help make intercourse last longer. (T)	94.1	36.5	57.6	(43.0, 72.2)
2. It's a good idea to use Vaseline for lubrication when using a condom. (F)	98.0	53.8	44.2	(30.1, 58.3)
3. If a woman is on the pill, her sex partners don't need to bother wearing a condom. (F)	98.0	61.5	36.5	(22.7, 50.3)
4. Some men never get symptoms for sexually transmitted disease. (T)	70.6	36.5	34.1	(16.0, 52.2)
5. For most sexually transmitted diseases, women usually get symptoms or early warning signs. (F)	84.3	51.9	32.4	(15.5, 49.3)
6. The condom should be unrolled before attempting to put it on the man's penis. (F)	88.2	61.5	26.7	(10.8, 42.6)
7. A condom should be worn so that it's snug at the tip. (F)	88.2	63.4	24.8	(9.0, 40.6)
8. The only way to make sure of preventing sexually transmitted disease is to use a condom every time. (T)	100.0	78.8	21.2	(10.1, 32.3)
9. When a person has more than one sex partner, it is impossible for that person to know for sure which partner was the source of the infection. (T)	90.2	71.2	19.0	(4.2, 33.8)
10. If a woman has a pelvic inflammatory disease (PID) more than once, she can become sterile and not be able to have children. (T)	98.0	80.8	17.2	(5.9, 28.5)
11. When a person has a sexually transmitted disease, that person should notify every sex partner he or she has had within the last month to get treatment. (T)	100.0	98.1	1.9	(−1.8, 5.6)
12. A person who has a sexually transmitted disease but no symptoms can still pass on that disease to other people. (T)	100.0	98.1	1.9	(−1.8, 5.6)
13. A man doesn't need to put on a condom until he's ready to come or ejaculate. (F)**	98.0	100.0	−2.0	(−5.8, 1.8)
Total Score (Estimated Mean)	11.1	7.9	3.2	(2.7, 3.7)

*Correct answers indicated in parentheses. T = True and F = False.
**This item was not included in the total score calculations.
NOTE: Intervention Group saw videotape; Control Group did not.

Differential Effectiveness of the Intervention

To determine whether the videotape was particularly helpful for subgroups of respondents, we estimated a series of regression models that tested the predictive power of each background characteristic as a main effect and in interaction with the intervention variable. Two sets of models were constructed—one for knowledge and one for attitudes.

When predicting knowledge, three background characteristics interacted with the intervention variable—years of education, previous condom use, and number of sexual partners. Despite correlations among these three background

TABLE 2—Percentage of Patients with Favorable Response on Each Item of the Attitude Index by Experimental Condition

Attitude Test Items*	% Favorable Responses			
	Intervention Group	Control Group	Difference	95% CI of Difference
1. Any man who refuses to use condoms is being selfish. (T)	94.1	61.5	32.6	(17.9, 47.3)
2. More women should get their lovers to use condoms. (T)	100.0	75.0	25.0	(13.2, 36.8)
3. There's no way a man and a woman can make using a condom sexy, even if they have the right attitude. (F)	86.3	63.5	22.8	(6.7, 38.9)
4. I can honestly say that, as a result of my visit to the clinic today, I will be using condoms more often. (T)	86.2	69.2	17.0	(1.3, 32.7)
5. I would not want my friends to know if I used condoms. (F)	90.2	75.0	15.2	(0.9, 29.5)
6. Condoms are not worth the bother. (F)	100.0	88.5	11.5	(2.8, 20.2)
7. A woman who insists that a man wear a condom is being unfair. (F)	96.1	84.6	11.5	(0.3, 22.7)
8. I just couldn't bring myself to talk to my sex partners about using condoms. (F)	94.1	82.7	11.4	(−0.7, 23.5)
9. I could never talk to my friends about whether I should use condoms. (F)	90.2	78.9	11.3	(−2.4, 25.0)
10. It's okay for a woman to carry condoms with her. (T)	94.1	88.5	5.6	(−5.2, 16.4)
11. People who use condoms every single time are being silly. (F)	100.0	96.2	3.8	(−1.4, 9.0)
12. If a friend tried to talk me into using condoms, I'd probably get angry. (F)	88.2	90.3	−2.1	(−14.0, 9.8)
13. Using a condom doesn't have to interfere with lovemaking. (T)	80.4	78.9	1.5	(−14.0, 17.0)
14. It is okay for a man to carry condoms with him. (T)	96.1	96.1	0.0	(−7.5, 7.5)
Total Score (Estimated Mean)	13.0	11.3	1.7	(1.0, 2.4)

*Answers indicating a favorable attitude are indicated in parentheses. T = True and F = False.
NOTE: Intervention Group saw videotape; Control group did not.

TABLE 3—Total Number of Strategies‡ Offered to Persuade Partners to Use Condoms by Experimental Condition

Number of Strategies	Experimental Condition	
	Intervention Group N (%)	Control Group N (%)
0	11 (21.6)	27 (51.9)
1	28 (54.9)	21 (40.4)
2	10 (19.6)	4 (7.7)
3	2 (3.9)	0 (0.0)
Mean*	1.1	0.6
Standard Deviation	0.6	0.4

*Difference between conditions is 0.5, 95% C.I. = 0.2, 0.8.
‡Exclusive of citing the condom's effectiveness in preventing pregnancy and disease.

characteristics, their effects persisted in a multiple regression model, as presented in Table 4 (total R^2=68.6 per cent). Regardless of background characteristics, subjects in the intervention group scored higher than subjects in the control group. Moreover, the videotape was most effective for subjects with less formal education, subjects who had never used condoms, and subjects with only one sexual partner in the past month.

When predicting attitudes, three background characteristics also interacted with the intervention variable—years of education, birthplace, and previous condom use. Because of correlations between previous use and the other two background characteristics, however, the effects of previous use (both main effects and interaction) did not persist in a multiple regression model. Thus, the final model, presented in Table 5, includes only the other two background characteristics (total R^2=44.3 per cent). Regardless of background characteristics, subjects in the intervention group had more favorable attitudes toward condoms than subjects in the control group. The intervention was most effective for subjects with less formal education and for those who were not native born.

STUDY 2: COUPON REDEMPTION

Method

The second study was conducted in the STD clinic at Boston City Hospital in the summer of 1986. As in the first study, clinicians referred all patients seen at the clinic, other than those who were under 18 years of age, were non-English speaking, had returned for a test-of-cure examination, or needed to be examined by a specialist. Of 393 patients referred, 57 were passed over to avoid creating a backlog of patients in the clinic; 32 were absent when called; five were non-English speaking; 25 were under 18 years of age; and 24 had participated previously in study 1 or an earlier pilot test.

Again, patients were seen while they awaited their laboratory test results. Of the 250 patients eligible to participate, 29 said they could not be reached by telephone for a follow-up interview we planned to conduct, and 39 refused to participate, leaving a study of 182 (or 72.8 per cent) of those eligible.

Qualifying patients who consented to participate were assigned at random to see the videotape (intervention group, n=89) or not (control group, n=93). Both groups of patients were given two postcard-sized coupons and told that they could be redeemed for free condoms; to protect patient privacy, the word "condom" did not appear on either coupon. The first coupon could be redeemed that day at another room at the clinic. The second, pre-addressed and stamped, could be redeemed by mail. When patients redeemed a coupon, they were given (or mailed) a packet with six condoms and a one-page information sheet on proper use.

Patient Profile

The subjects participating in this study were similar to

TABLE 5—Predicted Values of Attitude Index from Multiple Regression Model

Education	Birthplace	Attitude Index Scores		Difference
		Intervention Group	Control Group	
1 Junior High School	1 USA	12.9	9.4	3.5
	0 Other	12.1	8.6	3.5
2 High School	1 USA	13.1	11.1	2.0
	0 Other	12.3	10.3	2.0
3 Post High School	1 USA	13.4	12.8	0.6
	0 Other	12.6	12.0	0.6

Final Model
Attitude Index = 6.91 + 4.95 Intervention + 1.70 Education + .82 Birthplace − 1.47 (Education * Intervention)

TABLE 4—Predicted Values of Knowledge Test from Multiple Regression Model

Education	Previous Condom Use	Number of Partners in Past Month	Knowledge Test Scores		Difference
			Intervention Group	Control Group	
1 Junior High School	never	1	11.6	6.5	5.1
	never	5	10.4	6.7	3.7
	most times	1	10.7	7.4	3.3
	most times	5	9.5	7.6	1.9
2 High School	never	1	11.7	7.3	4.4
	never	5	10.5	7.5	3.0
	most times	1	10.8	8.2	2.6
	most times	5	9.6	8.4	1.2
3 Post High School	never	1	11.8	8.2	3.6
	never	5	10.6	8.4	2.2
	most times	1	10.9	9.1	1.8
	most times	5	9.7	9.3	0.4

Final Model
Knowledge Test = 5.37 + 6.70 Intervention + 0.81 Education + 0.30 Previous Use + 0.05 Number of Partners
−.70 (Education * Intervention) − .60 (Previous Use * Intervention)
−.35 (Number of Partners * Intervention)

those in Study 1: predominantly male (80.2 per cent), single (85.7 per cent), Black (85.2 per cent), and native born (74.7 per cent). Patients ranged in age between 18 and 73, with a median age of 24; 83.0 per cent of the patients were 30 years of age or younger. Again, educational attainment varied: 18.1 per cent had completed junior high, 43.4 per cent had completed high school, and 37.9 per cent had at least some post-high school education. Two-thirds of the subjects reported a prior history of sexually transmitted disease. Exactly half said they had two or more sex partners within the past month. There were no important differences between the two experimental conditions across all of these background variables.

Results

Coupon Redemption

Across both conditions, 110 (60.4 per cent) of the patients redeemed a coupon at the clinic only; another 25 patients (13.7 per cent) redeemed coupons both at the clinic and by mail; two patients (1.1 per cent) redeemed by mail only; only 45 subjects (24.7 per cent) did not redeem either coupon. Table 6 presents the number of coupons redeemed via clinic and/or mail by condition.

Patients who saw the videotape redeemed a significantly greater number of coupons; more control group subjects redeemed no coupons, while more intervention group subjects redeemed two. A Wilcoxon test yielded a p-value for the difference of .03. A logistic regression model showed that none of the background variables were associated with coupon redemption.

Discussion

Subjects who watched the videotape were more aware of their own susceptibility to STDs and their consequences, the benefits of condom use, the particulars of proper use, and general information on STD transmission than were subjects who did not see the videotape. Furthermore, intervention group subjects were more likely than control group subjects to recognize that one needs to use condoms to protect against disease, even if some other form of birth control is being used. Subjects who saw the videotape agreed more often that condom use could be "eroticized," that it is appropriate for women to insist that their lovers use condoms, and that men who refuse to us them are not thinking of their partner's well-being.

The intervention was most successful with those patients with fewer years of formal education. This finding has important clinical implications. In the various STD clinics in which we have worked, many clinicians have told us that they try harder to educate patients whom they assess to be smarter or better educated, often discounting larger numbers of patients as uneducable. The success of this intervention with those who are perceived as less educated speaks to the importance of not underestimating the potential of any patient to learn.

We feel that one of the reasons for the videotape's success was that it avoided a didactic, authoritarian tone, reflecting instead the language, beliefs, and values of the inner-city population to whom it was targeted.[35] However, just such cultural specificity is likely to render the videotape less effective with other cultural or socioeconomic groups. For example, campaigns to promote condom use among Hispanics, another high-risk group for AIDS, will have to take into account cultural discomfort with talking about sex in mixed groups of males and females and should use themes that are powerful in Latino culture, such as safeguarding fertility and protecting one's family.

Perhaps our most striking finding is the high proportion of patients from both intervention and control groups who sought free condoms by redeeming coupons. This finding strongly suggests that if STD clinics offer free condoms, they will be accepted by the vast majority of patients. Since the office where patients redeemed their coupons was immediately next door to the room where our research assistant offered the coupons, it is possible that the proximity of the two rooms, the relative ease with which coupons could be redeemed, and their packaging in unlabeled envelopes contributed to the high overall rate of redemption. Thus, while convenience factors were likely to have diminished the size of the difference between the groups, the high level of redemption in both groups speaks to the importance of engineering convenience into condom distribution programs. In addition, health educators may note that clinic redemption was much preferred to redemption by mail, a finding that seems to confirm social marketers' insistence on the importance of ease-of-purchase and accessibility to the product being promoted.[33,34]

The major limitation of this study is the difficulty in measuring changes in actual condom use: so far, researchers have been obliged to rely on self-reports and indirect proxy measures, such as coupon redemption. Methods for assessing actual condom use need to be developed and implemented.

Another limitation was our inability to track patients' knowledge and attitudes over time. Social science research points the importance of reaching the target audience repeatedly via multiple pathways.[2] As with any other one-time intervention, the effects of this videotape are unlikely to persist without reinforcement. The results of these studies suggest, however, that drama-based videotapes could be an effective component of an integrated program that continues over time.

The success with men of a videotape featuring a feminine perspective adds credibility to the observation made frequently by education intervention specialists that people are often more motivated to act out of concern for others than on their own behalf. There may be strategic advantages for condom promotion campaigns and other AIDS risk reduction efforts aimed at men to portray clearly the health risks that unprotected intercourse creates for women and children. Moreover, by portraying strong, assertive women taking responsibility for their health, educational programs have an opportunity not only to motivate men, but also to provide role models and social support to women who may wish to take more purposeful action on their own behalf. Ideally, inte-

TABLE 6—Total Number and Percentage of Coupons Redeemed by Experimental Condition

Number of Coupons Redeemed	Experimental Condition	
	Intervention Group *n* (%)	Control Group *n* (%)
0	17 (19.1)	28 (30.1)
1	56 (62.9)	56 (60.2)
2	16 (18.0)	9 (9.7)
Mean*	1.0	0.8

*Difference between conditions is 0.20, 95% CI = 0.02, 0.38

grated long-term programs will create a social climate in which both men and women will feel comfortable initiating condom use.

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Malpractice Seminar Report Available

A report on a recent medical malpractice seminar, co-sponsored by the Health Resources and Services Administration (HRSA) of the US Public Health Service and the Washington Business Group on Health, has been issued. Copies of the 47-page publication entitled Medical Malpractice: A Costly Barrier to Health Care System Reform may be obtained from the Washington Business Group on Health, 229 1/2 Pennsylvania Avenue, SE, Washington, DC 20003. Tel: (202) 547-6644.

Determinants of Needle Sharing among Intravenous Drug Users

Stephen Magura, PhD, Joel I. Grossman, PhD, Douglas S. Lipton, PhD, Qudsia Siddiqi, PhD, Janet Shapiro, MA, Ira Marion, and Kenneth R. Amann, MSW

Abstract: Data from 110 IV-drug abusing persons in methadone maintenance were analyzed to determine the correlates of needle sharing. Sharing was directly related to peer group behavior, attitudes conducive to sharing, economic motivation to share, not owning injection equipment, and fatalism about developing AIDS. Sharers were aware of their AIDS risk. Indicated measures to reduce needle sharing would be positive peer support groups to help resist pressures to share, legal and free access to fresh injection equipment, education on the utility of risk reduction, and increased treatment options for IV cocaine users. (*Am J Public Health* 1989; 79:459–462.)

Introduction

Intravenous drug users (IVDUs) are the largest group at risk for the acquired immunodeficiency syndrome (AIDS) in New York City.[1,2] IVDUs also are the major window for heterosexual transmission and main source of perinatal transmission of the disease.[3] The human immunodeficiency virus (HIV) can be transmitted through contaminated blood left in needles, syringes, and other injection equipment that may be shared. Although awareness and fear of AIDS reportedly have decreased sharing of injection paraphernalia, substantial numbers of addicts still engage in such behavior.[4–6] Entry into methadone treatment has been shown to reduce IV drug use,[7,8] but some individuals nevertheless continue or revert to drug injection and non-sterile needle sharing.[9,10]

Existing research on needle sharing is quite limited. Two ethnographic studies have identified various situational factors, economic considerations, and social pressures that may lead to sharing.[11,12] A survey of drug abusers in an inpatient treatment program found no demographic or personality differences between needle sharers and other drug abusers.[5] Both the amounts of AIDS knowledge and needle sharing among IVDUs appear to be high in the New York City area,[6,13,14] suggesting that knowledge may not affect sharing.

The present study examines addicts enrolled in methadone maintenance who continue to inject drugs, comparing those who share needles and other equipment with those who do not.

Methods

The study subjects are 110 persons in methadone maintenance who volunteered for an AIDS prevention demonstration research project and who also reported current IV drug use. The subjects are a subset of a larger sample of 274 volunteers, the remainder of whom did not report current IV drug use. Subjects were drawn from three methadone clinics in New York City between January and September 1987. The project sample is 27 per cent of the total patient population, and the study subjects (IVDUs) are 40 per cent of the project sample. Because project participants were self-selected, they may not be completely representative of the patient population.

From Narcotic and Drug Research, Inc. (Magura, Grossman, Lipton, Siddiqi, Shapiro), the Albert Einstein College of Medicine (Marion) and Long Island Jewish-Hillside Medical Center (Amann). Address reprint requests to Stephen Magura, Project Director, Narcotic and Drug Research, Inc., 11 Beach Street, New York, NY 10013. This paper, submitted to the Journal March 29, 1988, was revised and accepted for publication July 29, 1988.

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The data analyzed here were obtained from a self-administered questionnaire completed by all subjects at baseline. The questionnaire was not anonymous, but confidentiality was guaranteed; the results were not available to methadone clinic staff. Subjects were paid $15 for completing the questionnaire. The instrument consisted of fixed-response, Likert-type items written at an elementary reading level; it was also available in Spanish. Research staff were available on site to assist subjects who needed help with the questionnaires. Included were measures of knowledge about AIDS and its transmission; attitudes concerning AIDS prevention; and high-risk drug-using and sexual behaviors. IVDUs were classified as needle sharers if they reported any sharing of needles or other injection equipment within the previous month.

A set of independent variables for the analysis was derived from the questionnaire based on prior theory and research pertinent to needle-sharing behavior. These variables are: a) knowledge of AIDS risk-related behaviors; b) attitudes conducive to needle-sharing; c) intravenous drug use by one's peers; d) economic motivation to share needles; e) ownership of personal injection equipment; f) personal acquaintanceship with AIDS/ARC (AIDS-related complex) victims; g) perceived usefulness of avoiding AIDS risk behaviors; h) length of time enrolled in methadone treatment; and i) demographic characteristics. (Details of index construction are in the Appendix.)

The questionnaire also included data on needle sharers' attempts at risk-reduction, such as discretion in sharing, selectivity in partners, and equipment sterilization practices.

Results

Sociodemographic and drug use variables for the study sample are given in Table 1. Cocaine injection as well as cocaine combined with heroin ("speedballing") are predominant. The use of heroin with cocaine usually is intended to regulate cocaine effects.[15] Few subjects reported injecting heroin exclusively, thus indicating the effectiveness of methadone treatment in controlling heroin use. Methadone is not a substitute for cocaine. Forty per cent of the sample reported sharing needles or other injection equipment during the preceding month. About one-fourth of the subjects who shared needles reported doing so as part of visits to shooting gallaries; the remainder shared only in friendship networks or with sexual partners.

Table 2 presents the bivariate correlations between needle sharing and the independent variables. Some of the independent variables are also correlated among themselves (absolute mean intercorrelation = .23, range = .00 to .38). A multiple linear regression (OLS) analysis was conducted with

TABLE 1—Sociodemographic Characteristics, Drug Injection and Needle Sharing (N = 110)

Characteristics	Per Cent
Age (years)	
Under 30	17
30–34	42
35–39	22
40 and over	19
Gender	
Male	74
Female	26
Ethnicity	
White	64
Black	19
Hispanic	14
Other	3
Marital Status	
Single	72
Married	28
Employment	
Employed	43
Unemployed	57
Time in Methadone Program (years)	
Under 1	29
1–3	35
4 and over	36
Injected Cocaine in Last Month	75
"Speedballed" in Last Month	46
Shared Needles Equipment in Last Month	
Not at all	60
One or two times	24
One day a week	4
A couple days a week	8
Almost every day	4

the dependent variable coded as a dichotomy (some needle sharing in last month vs no sharing). Examination of residuals indicated that the data meet the linearity and homoscedasticity assumptions required for OLS.

The result of the regression analysis, which includes all

TABLE 2—Bivariate Associations between Independent Variables and Needle Sharing[a]

Variable	Correlation (r)	95% Confidence Interval[b]
Gender[c]	17	(−.03, .37)
Age	−.14	(−.33, .06)
Ethnicity[d]	−.04	(−.24, .16)
Marital Status[e]	23	(.04, .43)
Time in Methadone Program	19	(−.01, .39)
Knowledge of AIDS Risks[f]	02	(−.18, .22)
Attitudes Conducive to Sharing[g]	43	(.27, .66)
Peer IV Drug Use[h]	43	(.25, .64)
Economic Motivation to Share[i]	40	(.22, .62)
Does not Own Injection Equipment[j]	27	(.07, .47)
Knows AIDS ARC Victims[k]	09	(−.10, .29)
Perceived Utility of Risk Avoidance[l]	35	(.17, .57)

[a]Coded as no sharing = 0, some sharing = 1.
[b]To compute the confidence intervals, r was converted to the statistic Z (not standard score),[16] which has an approximately normal sampling distribution with standard error = $1 \sqrt{N-3}$. For our sample, SE = .10.
[c]Male = 0, female = 1.
[d]White = 0, non-White = 1.
[e]Married = 0, single = 1.
[f]Higher knowledge = higher score.
[g]More conducive attitudes = higher score.
[h]More peers using = higher score.
[i]More economic motives = higher score.
[j]Owns = 0, doesn't own = 1.
[k]More victims known = higher score.
[l]Low perceived utility = higher score.

TABLE 3—Multiple Linear Regression on Needle Sharing

Variable	Regression Coefficient	95% Confidence Interval
Gender	.101	(−.084, .287)
Age	.001	(−.014, .013)
Ethnicity	−.015	(−.179, .150)
Marital Status	.147	(−.026, .321)
Time in Methadone Program	.056	(−.128, .241)
Knowledge of AIDS Risks	.007	(−.025, .011)
Attitudes Conducive to Sharing	.098	(.022, .174)
Peer IV Drug Use	.093	(.010, .176)
Economic Motivation to Share	.132	(.016, .248)
Does not Own Injection Equipment	.252	(.039, .465)
Knows AIDS ARC Victims	−.030	(−.197, .137)
Perceived Utility of Risk Avoidance	.121	(.039, .204)
Intercept	−.711	

variables from Table 2, is shown in Table 3. Direct associations with needle sharing were found for attitudes conducive to sharing, peers' intravenous drug use, economic motivation to share, not owning personal injection equipment, and low perceived utility of AIDS risk avoidance. Knowledge of AIDS risks, knowing AIDS/ARC victims, gender, age, ethnicity, marital status and time in methadone treatment were not associated with sharing. Interaction effects among independent variables were not explored because of the large number of possible effects relative to the sample size.

Discussion

Apparently needle sharing cannot be attributed to ignorance about the consequences of such behavior, or how to effectively protect oneself and others against infection.

Attitudes toward sharing, however, do help explain this behavior. Addicts are more likely to share if they are unwilling to tolerate withdrawal symptoms in preference to sharing and/or if they believe that their friends would be "insulted" if they refused to share. Some can be induced to share when their friends claim that the needles are "clean." This may be due less to naivete than to reaching for an excuse to avoid threatened discomfort. Refusal to share after receiving assurances of safety may also signify distrust of one's peers.

Intravenous drug use by friends and sexual partners creates a social environment that leads to sharing. Thus our results support previous observations that peer behavior strongly influences needle sharing decisions.[11,12]

Economic pressure to share, as indicated by being unemployed and/or by a perceived inability to afford new needles, is directly associated with sharing. Limited funds often force addicts to choose between buying new injection equipment or additional drugs.

Not owning injection equipment is independently related to sharing, i.e., addicts who are reluctant to carry injection equipment, and therefore have less reason to own it, tend to rely on sharing.

The final variable identified through the regression analysis is Perceived Utility of Risk Avoidance—One interpretation of this result is that addicts with a fatalistic attitude, who believe that their past behavior has already made them highly vulnerable to developing AIDS, are prone to continuing needle sharing. An alternative interpretation is that needle sharers' fatalism is more of a reaction to their present

behavior, rather than a cause of that behavior. In that event needle sharers may simply be indicating their realistic awareness of their heightened risk for AIDS.

Needle sharers are not denying the dangers of their behavior and do report attempts to reduce the risks posed by sharing: most sharers say they have become more careful about those with whom they share needles and most claim they only borrow injection equipment in an "emergency" (Table 4). Although we cannot know how consistent or effective these risk-reduction efforts are, needle sharers are clearly aware of the implications of their conduct. Yet despite this knowledge, their desire to use a needle or to please their drug-using friends and sexual partners is strong and immediate.

Greater claimed selectivity in partners may be one reason that most needle sharers say they "never" or only "sometimes" sterilize their injection equipment after someone else uses it (Table 4). Another reason for infrequent sterilization may be that nearly half of sharers say that their friends would be "insulted if they sterilized the works" after others have used them. Finally, when injection equipment is "cleaned," it is often with ineffective sterilization techniques, e.g., water only.

The cross-sectional nature of this study makes it impossible to draw definitive conclusions about the presence or direction of causal effects. However, our statistical results are generally consistent with the findings of previous, mainly qualitative, studies. Assuming that the variables significantly associated with needle sharing in the regression model are indeed explanatory variables, some public health measures to help reduce this AIDS risk behavior can be recommended:

- One method to reduce needle-sharing would be to strengthen addicts' abilities to resist the leads of their peers, particularly (for women) of their sexual partners. What may help them are alternative sources of social support, e.g., peer groups or women's groups, where the interpersonal and communication skills necessary to resist pressures to share can be learned and reinforced.
- Free and legal access to needles and other injection equipment, not now existing in states with sizable IVDU populations, could be useful for addicts until they decide to enter treatment or begin to respond to treatment. This conclusion supports the concept of an experimental needle exchange that may soon be implemented in New York City.[17] Although this involves practical problems that are still not entirely resolved, experience with a needle exchange in Amsterdam has been favorable.[18]
- Because feelings of hopelessness or fatalism about developing AIDS may be inhibiting risk reduction, more emphasis should be placed on educating needle sharers about the potential benefits of behavioral change.
- Finally, much of the attraction of continued drug abuse among these subjects is attributable to intravenous cocaine. The data point to the need to increase community resources for preventing and treating cocaine abuse as one important weapon against the transmission of AIDS.

TABLE 4—Risk Reduction Attempts by Needle Sharers (N = 44)

	Per Cent
I've become a lot more careful who I shoot up with	77
I only share needles and works with certain friends	66
I only use somebody else's needles or works in an emergency	68
Cleans needles or works after someone else used	
Every or most times	29
Sometimes or never	71
Employs effective sterilization techniques*	50

*Soaks in bleach, pure alcohol, rubbing alcohol, or peroxide; or boils in water.

APPENDIX
Construction of Composite Measures

Knowledge of AIDS Risks

The correct answers were summed for 28 true-false questionnaire items (alpha = .83). Factor analysis of the knowledge item set did not reveal any separate subscales. Typical items are: "Anyone who has shot drugs might be carrying the AIDS virus now" (T); "If a needle looks clean it can't give you AIDS" (F); "You could get AIDS by having unprotected sex with a person who shot drugs" (T); and "Cleaning needles and works with soap and water will kill the AIDS virus" (F). (The full scale is available from the authors.)

Attitudes Conducive to Needle Sharing

This scale was constructed by counting responses to four items (alpha = .57): "I'd rather get sick than share works" (agree = 1); "My buddies would be insulted if I refused to share my works with them" (agree = 1); "If someone says their works are clean, I'll share" (agree = 1); and "We don't talk much about AIDS in my crowd" (agree = 1). (In the drug use context, addicts understand the term "sick" as "dope sick.")

Perceived Utility of Risk Avoidance

This scale sums responses to three items (alpha = .70): "I've already done plenty that could have exposed me to AIDS" (agree = 1); "My gut feeling is, I'm going to get AIDS sooner or later" (agree = 1); and "I never did anything that could give me AIDS" (disagree = 1).

Peer IV Drug Use

This index sums the scored responses to two items: "Some of my friends shoot up regularly" (agree = 2, not sure = 1, disagree = 0) and "In the last month, how many of your sexual partners were shooting drugs?" (two or more = 2, one = 1, none = 0).

Economic Motivation to Share

This index sums responses to two items: employment status (unemployed = 1) and "I can't afford to buy new needles all the time" (agree = 1).

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$4.4 Million in Nursing Grants Available in Fiscal 1989

Approximately $4.4 million is available in fiscal 1989 to support new grants for nurse practitioner and nurse midwifery programs, according to an announcement in the January 6 *Federal Register*.

The grants will assist in meeting the costs of projects to plan, develop, operate, expand or maintain educational programs. It is estimated that 23 competing projects averaging $185,000 will be supported.

Public or nonprofit private schools of nursing, schools of public health, hospitals, and other public or nonprofit private entities are eligible to apply for grants, as are medical schools that received a grant or contract under section 822(a) of the Public Health Service Act prior to October 1, 1985.

Applicants that demonstrate an increase in minority enrollment which exceeds the program's prior three-year average would receive a funding priority, according to the published notice.

The deadline for applications is July 1. Application materials are available from:

Grants Management Officer
Bureau of Health Professions
5600 Fishers Lane
Room 8C-22
Rockville, MD 20857
Tel: (301) 443-6960

Outbreaks of Hepatitis A among Illicit Drug Users, Oklahoma, 1984–87

JOHN HARKESS, MD, BARBARA GILDON, RN, AND GREGORY R. ISTRE, MD

Abstract: Transmission of hepatitis A has been described in many settings, although few outbreaks have been reported among illicit drug users. We report six unrelated outbreaks of hepatitis A among users of marijuana and intravenously administered methamphetamine. Although the exact mode of transmission could not be determined, practices associated with illicit drug use may have facilitated transmission of hepatitis A. These outbreaks and other recent information indicate that illicit drug users may be at increased risk of acquiring hepatitis A infection. (*Am J Public Health* 1989; 79: 463–466.)

Introduction

Outbreaks of hepatitis A infection have resulted from common-source food and water-borne transmission and from person to person transmission in households, day care centers, institutions for the mentally retarded, and in communities or other settings where sanitation is poor.[1–3] Although parenteral drug users are known to be at increased risk for many infections including hepatitis B, human immunodeficiency virus, and infective endocarditis,[1,4,5] they are not generally recognized to be at increased risk of hepatitis A.

Reports from Scandinavian countries have drawn attention to the occurrence of hepatitis A infections among parenteral drug abusers.[6–8] In two studies from Sweden, 42 per cent and 58 per cent of patients diagnosed with hepatitis A were found to be parenteral drug users[7,8] and prolonged outbreaks among drug users have been observed.[8,9] In the United States, 10 per cent of patients diagnosed with hepatitis A in five metropolitan counties reported using illicit drugs[10] and outbreaks of hepatitis A among illicit drug users have recently been identified in some areas.[11,12]

We report six unrelated outbreaks of hepatitis A infection that occurred among parenteral and non-parenteral drug users in Oklahoma.

Methods

A case of hepatitis A was defined as jaundice and a positive test for the IgM-specific antibody to hepatitis A virus (IgM-specific anti-HAV), or jaundice with at least two other symptoms of hepatitis (anorexia, nausea, vomiting, fever, dark urine) and no other explanation for the illness. A primary case-patient was defined as a person with hepatitis A who had not lived in the same household or had sexual contact with anyone having hepatitis A in the two to six weeks before onset of illness. A secondary case-patient was defined as a person with hepatitis A who had household or sexual contact with another person with hepatitis A in the two to six weeks before onset of illness.

Outbreaks were identified by communicable disease nurses through routine disease surveillance. In each outbreak investigation, additional hepatitis A case reports were solicited by telephone from physicians and hospital laboratories. Patients were interviewed in person about possible exposures in the two to six weeks before onset of illness including contact with any person having symptoms of hepatitis, social events attended, restaurants patronized, sources of drinking water, contact with a child in a day care center, and use of illicit drugs. For three outbreaks, age- (±5 years), sex-, and neighborhood-matched controls were identified by door-to-door survey using predetermined search patterns and interviewed in the same manner as the case-patients. Confidence intervals for odds ratios were calculated using an exact method.[13]

Results

Outbreak 1

Between October 1984 and March 1985, 13 cases of hepatitis A were reported in County A, Oklahoma compared to only a single case in the previous nine months. Investigation identified 45 persons with possible hepatitis. Illness in 29 persons met the case definition for hepatitis A: 18 persons had jaundice and a positive IgM-specific anti-HAV, and 11 persons who were not tested had illness which met the clinical case definition.

The 29 case-patients were White, ranging in age from 12–48 years (median = 24 years). 22 (76 per cent) were male. Illness onsets occurred between October 1984 and March 1985 (Figure 1A). Of the 29 case-patients, 27 (93 per cent) were fully interviewed. No common exposures to any food, eating establishment, or social gathering were identified, nor did any patients have household contact with a child attending a day care center in the two to six weeks preceding illness.

Of the 27 fully interviewed case-patients, 20 (74 per cent) had smoked marijuana and 13 (48 per cent) admitted to intravenous drug use in the two to six weeks before illness. Of the 13 IV drug users, all had injected illegally manufactured methamphetamine and eight had also injected cocaine. In contrast, of the 16 persons who were interviewed but whose illnesses did not meet the hepatitis A case definition, only one smoked marijuana during the two to six weeks before illness (odds ratio [OR] = 42.9, 95% confidence interval [CI] = 4.5-1891), and none had injected illicit drugs (95% CI lower bound = 3.24).

Of the 27 case-patients, 17 had not had household or sexual contact with a person having hepatitis A during the two to six week period preceding their illness and were classified as primary case-patients; 10 patients who had sexual or household contact with another case-patient were classified as secondary case-patients. Of the 17 primary case-patients, 15 (88 per cent) had used marijuana during the two to six weeks before illness compared to only five (50 per cent) of 10 secondary case-patients (OR = 7.5, 95% CI = 0.8-94.1). Additionally, 10 (59 per cent) of 17 primary case-patients admitted to injecting an illicit drug compared to three (30 per cent) of 10 secondary case-patients (OR = 3.3, 95% CI = 0.5-26.1).

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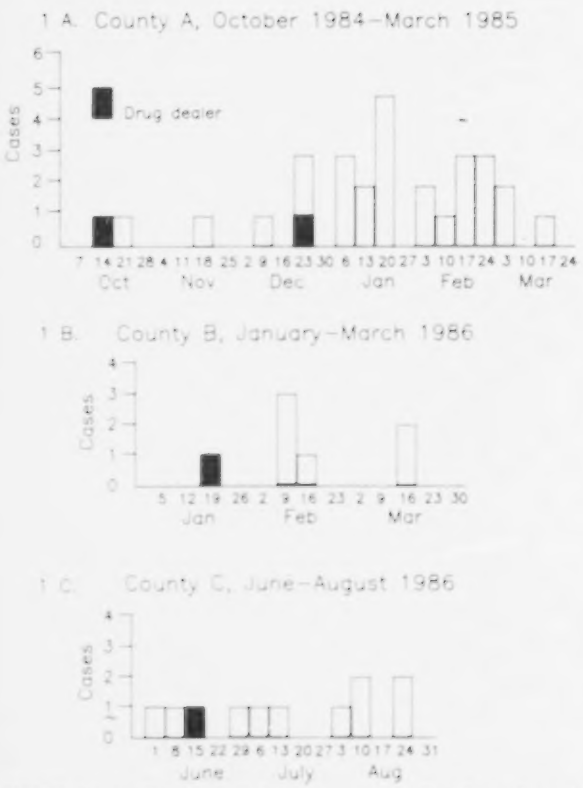
1 A. County A, October 1984–March 1985
Cases
Drug dealer
7 14 21 28 4 11 18 25 2 9 16 23 30 6 13 20 27 3 10 17 24 3 10 17 24
Oct Nov Dec Jan Feb Mar
1 B. County B, January–March 1986
Cases
5 12 19 26 2 9 16 23 2 9 16 23 30
Jan Feb Mar
1 C. County C, June–August 1986
Cases
1 8 15 22 29 6 13 20 27 3 10 17 24 31
June July Aug

FIGURE 1 A–C.—Hepatitis A Outbreaks among Illicit Drug Users, Cases by Week of Onset, Oklahoma, 1984–87

Matched controls were found for 12 of the 17 primary case-patients. Case-patients were similar to controls with respect to race (all White), employment (62 per cent of case-patients vs 77 per cent of controls employed), and exposure to a child in a day care center (0 vs 8 per cent); however, case-patients were less likely to be married (15 vs 77 per cent) than were controls. All 12 primary case-patients had used marijuana compared to only three of 13 (23 per cent) controls (95% CI lower bound = 2.5), and 10 of 12 primary case-patients had injected drugs compared to none of the controls (95% CI lower bound = 2.9).

Outbreak 2

Between January and March 1986, seven patients with hepatitis A were identified in County B, Oklahoma (Figure 1B). No cases had been reported in that county for the previous six months. The patients ranged in age from 19 to 33 years (median = 26.5), five (71 per cent) were male, and all were White. Six (86 per cent) patients had a positive IgM-specific anti-HAV; one refused phlebotomy.

On investigation, no common social gathering or eating establishment was identified, and no case-patient had had household contact with a child in a day care center. One patient who had been a household contact of a jaundiced person during his exposure period was classified as a secondary case. None of the sexual or household contacts identified by the remaining six case-patients were known to have had clinical hepatitis during the patient's exposure period.

All six primary case-patients had used an illicit drug during their incubation periods; five regularly smoked marijuana and five had intravenously injected methamphetamine. For three case-patients, methamphetamine had been prepared and injected for them by someone who was ill with hepatitis A during the two to six weeks before illness. A fourth patient had shared a marijuana cigarette with a jaundiced person in the two to six weeks before his onset.

Two matched controls each were found for five of the six primary case-patients. Four (80 per cent) of five primary case-patients had smoked marijuana compared to two (20 per cent) of 10 controls (OR = 7.0, 95% CI = 0.6, 334), and four (80 per cent) primary case-patients had injected drugs compared to none of the controls (95% CI lower bound = 1.8).

Outbreak 3

Between June 1 and August 29, 1986, 12 persons became ill with hepatitis A in County C, Oklahoma (Figure 1C) compared to only four cases reported during the preceding 17 months. One patient was 3 years old, the remainder ranged in age from 14 to 27 years (median = 23). Eight (67 per cent) were male, and all were White. Eleven (92 per cent) patients had a positive IgM-specific anti-HAV; one refused phlebotomy.

No common exposure to an eating establishment, social gathering, or day care center could be identified. Of the 12 case-patients, 11 (92 per cent) were classified as primary case-patients and one was classified as a secondary case-patient. None of the eight sexual contacts identified by seven of the 11 primary case-patients were homosexual men and none had clinical hepatitis in the preceding months.

Of the 11 primary case-patients, eight (73 per cent) had smoked marijuana, two (18 per cent) had consumed raw hallucinogen-containing mushrooms, and four (36 per cent) had injected illegally made methamphetamine during the two to six weeks before illness. Matched controls were found for 10 of the 11 primary case-patients; no control was sought for the 3-year-old case-patient. Of the 10 primary case-patients, eight (80 per cent) had used an illicit drug during the two to six week incubation period compared to none of the control patients (95% CI lower bound = 2.2). Additionally, three case-patients had interacted with someone who had hepatitis during the two to six weeks before onset of illness, but in all three cases the contact did not include sex, consumption of food prepared by the ill person, or living in the same household. One case-patient had consumed raw hallucinogen-containing mushrooms that were purchased from another ill case-patient, one patient had shared a marijuana cigarette while visiting with another ill case-patient, and one patient had shared a marijuana cigarette and intravenously injected amphetamines purchased from another ill case-patient.

Outbreaks 4–6

Three additional outbreaks of hepatitis A among illicit drug users have been investigated in Oklahoma since 1984. Nine cases occurred in County D from September to December 1985 (Figure 1D), six cases in Town E from November 1986 to March 1987 (Figure 1E), and 10 cases in County F from October to December 1987 (Figure 1F). As in the outbreaks already described, most cases occurred among young adults males over a period of two to five months (Table 1). In five of the six outbreaks, one or more of the case-patients was identified as a dealer or distributor of illicit drugs. In all five instances, cases among dealers occurred early in the outbreak and before illnesses in persons known to have been customers (Figures 1A–F).

TABLE 1—Summary Data for Six Unrelated Outbreaks of Hepatitis A among Illicit Drug Users, Oklahoma, 1984–87

Location	Duration (months)	Total Cases	Age Range (years)	% 15–25 Years of Age	% Males	Smoked Marijuana			Injected Drugs			Used Any Illicit Drug		
						%	No	Total*	%	No	Total*	%	No	Total*
County A	5	29	12–48	83	76	74	20	27	48	13	27	74	20	27
County B	3	7	19–33	100	71	83	5	7	83	5	7	86	6	7
County C	3	12	3–27	83	67	73	8	11	36	4	11	73	8	11
County D	3	10	11–39	70	80	86	6	7	50	4	8	75	6	8
Town E	5	11	16–35	100	91	100	11	11	73	8	11	100	11	11
County F	2	10	20–35	100	90	100	10	10	80	8	10	100	10	10

*Total for whom information was available

Discussion

In the hepatitis A outbreaks reported here, illness occurred almost exclusively among young adult men, most of whom had used illicit drugs during their time of likely infection with hepatitis A virus. No evidence of common-source transmission was found, and most patients had not had close personal contact with another person with hepatitis. Findings which suggested that drug use practices may have played a role in transmission of these infections include:

- In Outbreak 1, marijuana and parenteral drug use was more common among case-patients than among persons whose illnesses did not meet the case definition;
- In Outbreak 1, illicit drug use was more common among primary case-patients whose source of infection was uncertain than among secondary case-patients whose likely sources of infection were apparent;
- Primary case-patients in three outbreaks were more likely than matched neighborhood controls to have used illicit drugs;
- the most likely sources of infection for seven patients in Outbreaks 2 and 3 were other ill persons who had sold, shared, handled, or prepared and injected drugs for them. As best we could determine, other contact between these persons was casual and did not involve sexual contact, living in the same household, or the consumption of food prepared by the ill person.
- Illnesses among drug dealers, who handled, sold, and helped others to inject drugs, occurred early in the outbreaks and preceded illnesses among known customers.

Poor personal hygiene and inferior living conditions which may accompany illicit drug use have been advanced as possible explanations for hepatitis A outbreaks among drug users in Scandinavia.[6] However, it is possible that drugs and associated apparatus may be vehicles for transmission of hepatitis A. The conditions under which illicit drugs are produced, transported, mixed with other substances, and prepared for use may allow fecal contamination to occur. Fecal contamination of rectally carried drugs was thought to have been responsible for one outbreak of hepatitis A among intravenous drug users in Sweden.[9] In the outbreaks reported here, persons ill with hepatitis A may have fecally contaminated spoons, cotton filters, or the drug itself while preparing and intravenously administering it to persons who subsequently became ill. Transmission of hepatitis A by blood transfusion has occurred[14,15] so parenteral transmission through the use of contaminated needles may also be possible.

Some case-patients admitted only to having smoked marijuana. Patients we interviewed described frequently handling marijuana while rolling cigarettes and occasionally swallowing small particles left in their mouths after inhaling from hand-rolled cigarettes. Contamination of marijuana with fecal coliforms was demonstrated in samples obtained during an outbreak of hepatitis A among drug users in Alaska,[11] and smoking or handling of fecally contaminated marijuana has been implicated in the transmission of salmonellosis.[16] Since

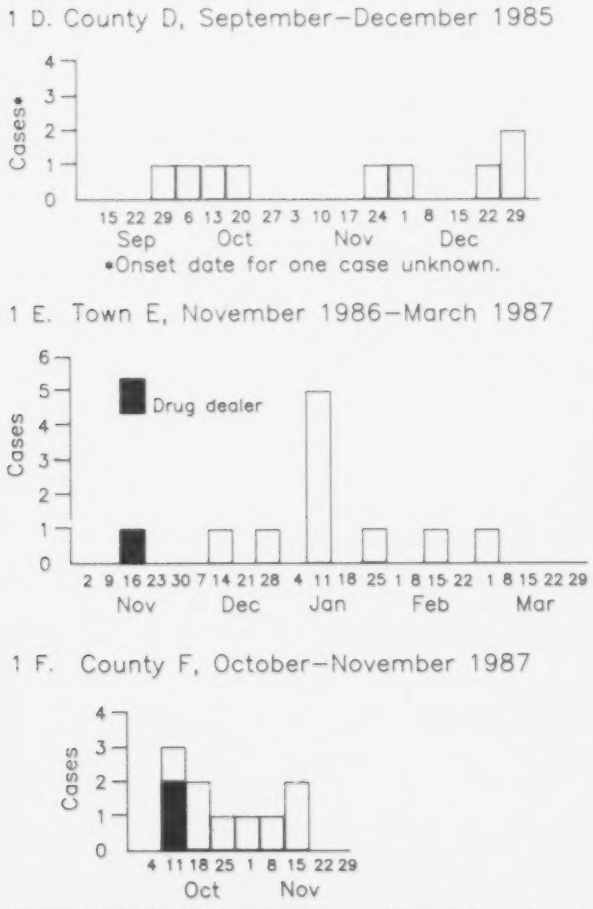

FIGURE 2 D–F—Hepatitis A Outbreaks among Illicit Drug Users, Cases by Week of Onset, Oklahoma, 1985–87

hepatitis A virus has been shown to survive as long as a month in dried feces,[17] it is possible that fecally contaminated marijuana could be a vehicle for hepatitis A.

The small number of cases that occurred in these outbreaks and the difficulties in obtaining accurate information from patients regarding the use and origin of illicit drugs make it difficult to determine the exact mode of transmission. With the imminent development of a vaccine for hepatitis A,[18] the identification of illicit drug users as being at increased risk of infection may have implications for prevention.

ACKNOWLEDGMENTS

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New AIDS Prevention Research Center to be Established at Columbia University

Columbia University is creating a $3.6 million center in its School of Social Work that will study and help apply measures to prevent the spread of AIDS among young people in minority groups. The new AIDS Prevention Research Center will develop and implement a variety of community-based techniques in the New York area to prevent the spread of AIDS among minority youths through the reduction of intravenous drug use and unsafe sexual behavior. The center will focus its efforts in Harlem and Northern Manhattan.

Minority youths are being targeted because the incidence of AIDS among minority groups is disproportionately high; Blacks and Hispanics accounted for 40 per cent of AIDS cases nationwide as of October.

The center will be supported by a five-year, $3.6 million grant from the National Institute on Drug Abuse, obtained with the support of Rep. Charles B. Rangel of New York.

Program participants, 8 to 16 years old, will be taught how to use problem-solving skills and to develop alternatives to high-risk activities. Assertiveness training, avoiding drug use, and learning safe sexual practices will be covered. Participants will not be tested for AIDS. Through personal and computer instruction, they will be taught ways to avoid exposure to AIDS. Information will be tailored to the youths' particular neighborhoods.

In supporting the center, Rep. Rangel said, "The survival of the minority community is dependent upon our arresting the AIDS epidemic, which means stopping the drug abuse that is responsible for so many of these cases. The new AIDS Prevention Research Center provides the opportunity for scientists and researchers to become intimately involved with the communities most devastated by the AIDS epidemic and equip members of those communities with the skills necessary so that they, themselves, can continue to fight this epidemic in the months and years to come."

Race and the Prevalence of Syphilis Seroreactivity in the United States Population: A National Sero-Epidemiologic Study

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Abstract: We used the 1978 National Health and Nutrition Examination Survey to examine the prevalence of positive syphilis serologies in the US population. Analysis of risk markers—gender, age, marital status, education, income, and residence—indicates that all except gender are associated with syphilis seroreactivity, independent of race. Controlling for associated risk markers, the Black-White odds ratio of syphilis seroreactivity is 4.7 (95% CI=2.7, 8.2). Current knowledge of racial differences in sexual and health care behavior does not explain the Black-White difference in the prevalence of syphilis seroreactivity. (*Am J Public Health* 1989; 79:467–470.)

Introduction

During the past 50 years, surveys with varying design have reported large racial differences in the incidence and prevalence of syphilis seroreactivity in the United States.[1–6] Among four million young male military recruits in the 1940s, syphilis rates were more than 13 times higher in Blacks than in Whites[1,2]; in community studies in Baltimore in 1939[4] and Savannah in 1945,[5] Black males were seroreactive at rates 9 to 16 times the rates for White males, and Black females between 16 and 21 times those of White females. Similarly, the first US population-based National Health Examination Survey (NHES-I) in 1960–62 found evidence of syphilis seroreactivity in 10 times as many Black as White men, and in eight times as many Black as White women.[6] Finally, a study of reported primary and secondary syphilis in New York State (outside of New York City) between 1975 and 1981 indicated a Black-White ratio of 8.4.[6]

To reduce the excess risk of syphilis among Blacks, and to control the incidence of syphilis in the general public, we must know not only the distribution of risk in the population, but also those aspects of the social environment which are related to incidence and which may be altered to reduce it. However, with the exception of NHES-I,[6] previous epidemiological analyses of racial differences[2–4] have been limited by sample selection and size, or by restriction to a small segment of the US population.

To further explore these relationships, we examined the distribution of positive syphilis serology in the US by analysis of the second (1978) National Health and Nutrition Examination Survey (NHANES-II)—the largest population-based health examination survey of the US conducted to date.[7] We tested the hypothesis that racial differences in syphilis seroreactivity are explained by differences in the distribution of sociodemographic risk markers—gender, age, marital status, residence, income, and education.

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Methods

Between 1976 and 1980, the National Center for Health Statistics conducted a stratified probability cluster survey of the US population, NHANES-II.[7] Of 27,801 civilian, noninstitutionalized individuals selected, 20,243 were Whites and Blacks 12 years of age and older. In this study we considered only Blacks and Whites age 12 and older. Of these, 18,094 (89 per cent) completed the interview regarding demographics and health and nutritional status; 14,196 subjects (70 per cent of those selected) were administered physical and laboratory examinations. The interview did not include questions on sexual history or history of sexually transmitted diseases (STD); the physical examination did not include the genitalia. Sera sufficient for syphilis testing were available from 12,989 subjects, 92 per cent of those examined. Tests for syphilis seroreactivity were conducted by the Sexually Transmitted Disease Laboratory Program at the Centers for Disease Control.

Sera were screened with the Automated Reagin Test (ART) and quantified with the Rapid Plasma Reagin Test (RPR). Sera reactive in the screening test were confirmed with subsequent reactive results on either the Microhemagglutination Assay for Treponema Pallidum (MHA-TP) or the Fluorescent Treponemal Antibody Absorption Test (FTA-ABS).[8,9]

Projections of US population characteristics based on the NHANES-II sample examined are generally comparable to projections based on the total sample.[10] Similarly, comparison of projected population characteristics from the NHANES-II sample interviewed but not tested for syphilis with projections from the sample tested suggests no major bias in study estimates based on the sample tested (see Appendix).

In our estimation of prevalences and odds ratios of seroreactivity, each subject was assigned a weight based on three characteristics: the probability of being sampled by NHANES-II, refusal to be examined (adjusted for age, income, urban residence, and region), and a poststratification factor to ensure that the weighted sample agreed with that of the US population in mid-1978 in distributions of gender, race, and age. Standard errors of prevalence and odds ratio estimates were made with a first order Taylor series approximation, incorporating the complex sample design.[11,12]

Several sociodemographic risk markers are associated with race in the NHANES-II population. Blacks are younger and less likely to be married; they live more commonly in central cities, have less education and lower incomes. These sociodemographic risk markers might confound the association of race and syphilis seroreactivity.

We estimated odds ratios for each sociodemographic risk marker (i.e., race, gender, age, income, etc.) while controlling for other factors, using logistic regression.[11] Because eventual education and marital status are less likely to be established by age 20, we restricted these analyses to subjects age 20 years and older. We also ascertained odds ratios for Blacks versus Whites in substrata defined simultaneously by education, income, and marital status, for which the sample was large enough ($n \geq 50$ for each race) for reasonable estimation.

Results

The prevalence of syphilis seroreactivity among persons 12 years of age and older in the 1978 US population is estimated to be 0.81 per cent (Table 1). The prevalence among Blacks is 3.05 per cent and among Whites it is 0.53 per cent, for a crude prevalence ratio of 5.7 (95% CI = 3.8, 8.6).

As expected, the prevalence of seroreactivity is related to age, marital status, residence, income, and education; prevalence is little affected by gender (Table 1). For the most part, the associations of these sociodemographic risk markers and syphilis seroreactivity hold both for the population as a whole and for the population stratified by race (Table 1). Prevalences for seroreactivity in Blacks relative to Whites are substantially elevated in all sociodemographic strata, except for those with income greater than $15,000.

The odds ratio of positive serologies, Blacks versus Whites, is 4.7 (95% CI = 2.7, 8.2) when the effects of other variables are controlled by logistic regression (Table 2). There is no statistically significant interaction between race and any of the variables for which we controlled. Prevalence of seroreactivity increases with age, and decreases with college education and greater income. The higher prevalences of seroreactivity among non-married persons and persons living in large cities (SMSAs) are reduced substantially after controlling for other risk markers. The independence of the association between race and seroreactivity is further supported by stratified analysis (data available on request to author).

TABLE 2—Odds Ratios of Syphilis Seroreactivity for Each Sociodemographic Risk Marker, Independent of all Others

Risk Marker	Odds Ratio (95% CI)
Race (Black vs White)	4.74 (2.74–8.20)
Age (per year)	1.05 (1.04–1.06)
Income (per $5,000 decrement)	1.30 (1.01–1.67)
Education	
No High School (HS) vs HS	1.14 (0.62–2.09)
HS vs some college	2.48 (1.07–5.76)
Residence (large city vs not)	1.30 (0.77–2.19)
Marital Status (not married vs married)	1.47 (0.82–2.66)
Sex (male vs female)	1.27 (0.77–2.10)

Discussion

We found Black race to be a strong risk marker of syphilis seroreactivity in the US population, independent of other standard sociodemographic risk markers, a finding consistent with earlier studies. While previous studies have used different serological tests and test sequences, our findings, along with another recent investigation,[6] indicate that the race differential may be narrowing. The crude prevalence ratios of syphilis for Blacks versus Whites appear to have declined from between 9 and 21 in the 1940s, to 8–10 in the 1960s, to 5–8 in the late 1970s. Nevertheless, a large and significant racial difference persists.

Because the syphilis tests used in NHANES-II may become nonreactive after longstanding infection, with or without treatment,[13] or after treatment of early syphilis,[14] this sequence provides a conservative estimate of prior

TABLE 1—Prevalence of Syphilis Seroreactivity in NHANES-II: Per Cent Seropositive (± 1.96 se)

	Total	White	Black	Relative Prevalences (95% CI)
Overall	0.31 (± .17)	0.53 (± .14)	3.05 (± .87)	5.7 (3.8,8.6)
Sex				
Male	.79 (± .27)	.57 (± .24)	2.68 (± .90)	4.7 (3.0,7.2)
Female	.83 (± .20)	.50 (± .18)	3.36 (±1.33)	6.8 (3.7,12.4)
Age (years)				
12–29	.10 (± .07)	.06 (± .06)	.39 (± .44)	6.2 (1.2,33.5)
30–44	.67 (± .31)	.42 (± .28)	2.88 (±2.23)	6.9 (2.2,21.1)
45–59	1.05 (± .41)	.60 (± .42)	5.08 (±2.69)	8.5 (3.1,23.4)
60–74	2.75 (± .63)	1.92 (± .55)	12.28 (±2.78)	6.4 (4.5,9.0)
Education*				
<High School	2.13 (± .59)	1.27 (± .41)	6.52 (±2.12)	5.1 (3.4,7.8)
High School	.77 (± .31)	.57 (± .35)	2.88 (±1.63)	5.1 (2.0,12.8)
>High School	.33 (± .16)	.28 (± .18)	1.01 (±1.09)	3.5 (0.8,14.9)
Income				
<$6,000	2.42 (± .83)	1.45 (± .67)	5.91 (±1.92)	4.1 (2.5,6.7)
$6–9,999	1.03 (± .51)	.68 (± .33)	3.02 (±2.45)	4.4 (1.8,10.8)
$10–15,000	.54 (± .19)	.42 (± .25)	1.79 (±2.10)	4.3 (0.8,21.8)
>$15,000	.28 (± .17)	.28 (± .17)	.32 (± .13)	1.2 (0.6,2.3)
Marital Status*				
Single	0.63 (±0.45)	0.46 (±0.46)	1.59 (±1.10)	3.5 (1.2,10.5)
Married	0.72 (±0.21)	0.48 (±0.16)	3.51 (±1.67)	7.3 (4.1,12.8)
Separated Divorced	2.16 (±0.96)	1.14 (±0.60)	6.16 (±3.94)	5.4 (2.3,12.3)
Widowed	3.28 (±1.59)	2.22 (±1.37)	9.07 (±4.03)	4.1 (2.2,7.6)
Residence				
SMSA, Central City	1.39 (± .55)	.82 (± .43)	3.12 (±1.08)	3.8 (2.5,5.7)
SMSA, Non-Central City	.53 (± .24)	.43 (± .25)	2.14 (±1.74)	4.9 (1.6,15.0)
Non-SMSA	.64 (± .27)	.46 (± .24)	3.72 (±2.49)	8.2 (3.5,19.0)

*only those over 20 are considered
**Standard Metropolitan Sampling Area

infection. The test sequence also introduces a possible bias in the comparison of race-specific seroreactivity. If Blacks receive medical treatment less frequently, or delay more often than Whites in seeking treatment, they would be more likely than Whites to be detected by the NHANES-II test sequence, given infection. Such ascertainment bias would artificially raise the odds ratios of Blacks over Whites.

Black-White differences in syphilis seroreactivity might be mediated by increased susceptibility to infection among Blacks, given the same exposure, or by greater capacity to sustain or transmit infection. However, while Blacks and Whites are reported to manifest differing symptoms in later stages of syphilis,[15] no studies have demonstrated either differential susceptibility to infection or differential capacity to transmit. Studies which purport to show racial differences in the likelihood of disease transmission per sexual contact[16] have not assured comparable exposures among Black and White subjects.

Higher rates among Blacks than among Whites have been found for a variety of sexually transmitted diseases. Reported gonorrhea is 10 times higher in Blacks in the US population[17*] and between 4.5 and 10.7 higher among women of reproductive ages, adjusted for several measures of sexual activity.[18] Black women in the US population report 1.8 times the rate of ambulatory or hospitalized treatment for pelvic inflammatory disease (PID) as do Whites,[19] and observed hospitalization for PID is 2.5 times more common in Blacks than in Whites.[20] Rates of infertility are 1.9 times greater in Blacks than in Whites.[21] Herpes simplex virus type 2 antibody, controlling for gender, is 3.4 times higher in Blacks,[**] while hepatitis B antibody, controlling for a variety of sociodemographic risk markers, is 4.6 times higher in Blacks.[†] AIDS has affected Blacks at a rate 3.1 times that for Whites.[22] Cervical cancer, suspected of an STD etiology,[23] is 2.3 times more common in Blacks.[24]

Population-based surveys of the US[25] have not yet provided comprehensive information on sexual behavior or racial differences in sexual behavior. The most current information compares White and Black women of reproductive age in Los Angeles[‡] and in the US population.[26] Several reported characteristics of Black women place them at greater risk of STD. The initiate sexual activity earlier and are more sexually active until age 25, when this difference diminishes[26,‡]; they use barrier methods of contraception less frequently than Whites.[‡] Both Black and White women prefer same ethnicity sexual partners[‡]; thus Blacks are more likely than Whites to have sex in a high prevalence population.

In contrast, other characteristics place Whites at greater risk of STD. Whites report greater numbers of sexual partners than Blacks,[‡] and more frequent sexual activity.[‡,26] From these initial current findings, the overall picture of (female) Black-White differences in sexual behavior does not clearly put either group at higher risk.

Evidence on racial differences in treatment for syphilis, other STDs, and other health problems is also sketchy and inconsistent. With regard to health care in general, Blacks are far less likely to be covered by private or federal health insurance[27]; in both good and poor health, Blacks have consistently fewer physician contacts per year.[28] With regard to treatment for STDs, in Columbus, Ohio, fewer White than Black heterosexual patients reported curtailing their sexual activity when they perceive dysuria or urinary tract discharge; patients' response to symptoms of syphilis were not elicited in this study.[29] Whites also delay clinic visits between 30 per cent and 90 per cent longer than Blacks with similar symptoms.[29] It is unclear how findings on health care seeking in general or of sexually transmitted disease clinic patients with discharge or dysuria might apply to patients from the general population with syphilis infection.

Race specific differences may be explained in part by historical differences in prevalence and by the predominance of intraracial sexual relations. While the present study demonstrates the significantly greater prevalence of syphilis seroreactivity among Blacks than among Whites, we do not yet know enough about racial differences in sexual behavior through which infection is acquired, or about racial differences in ways in which infection is eliminated to explain persisting differences in prevalence.

APPENDIX
Demographic Characteristics of Population Projections from NHANES-II Sample Not-tested and Tested

Characteristics	Not-Tested	Tested
Sex		
Male	45.6	47.8
Female	54.5	52.2
Race		
White	86.4	86.7
Black	11.6	10.7
Other	2.0	2.6
Age (years)		
12–29	37.5	41.9
30–44	22.4	23.7
45–59	22.3	20.2
60–74	17.8	14.3
Education		
< HS	32.1	35.9
HS	33.7	28.5
Some college	34.2	35.7
Residence		
Central City	30.9	27.9
SMSA, non-central	34.1	34.7
Non-SMSA	35.1	37.3
Income ($)		
< 6,000	15.8	14.2
6–9,999	21.9	20.3
10–15,000	17.1	18.2
< 15,000	39.0	43.5
unknown	6.3	3.9
Region		
Northeast	26.1	22.4
Midwest	24.5	25.1
South	26.9	25.6
West	22.6	27.0
Marital Status		
Married	67.7	69.6
Single	14.2	14.8
Separated Divorced	10.3	9.2
Widowed	7.5	6.1

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New Name: Advisory Council on Nurses Education

A name change—from the former National Advisory Council on Nurse Training—and a new charter were announced recently by HRSA. Now known as the Advisory Council on Nurses Education, the council will be increased in size from 19 to 21 members appointed by the Secretary of Health and Human Services. The changes implement provisions of the Nursing Shortage Reduction and Education Act of 1988.

The nursing council will continue to advise and make recommendations to the HHS Secretary and the HRSA Administrator on regulations and policy matters arising from the administration of nurse training programs, and review grant applications and make recommendations for various nurse education programs authorized under Title VIII of the Public Health Service Act.

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Mumps in the US Army 1980–86: Should Recruits Be Immunized?

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Abstract: The US Army's experience with mumps hospitalizations was examined for the years 1980 through 1986. One hundred fifty-two cases among active duty Army soldiers were identified. Mumps rates declined from 3.85 per 100,000 active duty soldiers per year in 1980 to 1.28 in 1985, but an outbreak during 1986 caused rates to jump to 6.65. Attack rates were found to decline dramatically with increasing age or length of military service, with 74 per cent of cases occurring in soldiers with three years or less of service. Rates for Blacks and Whites were similar, but were higher for other minorities. Complications reported were mild. A cost-benefit analysis, assuming all recruits were to be vaccinated, estimated average annual vaccination program costs of $286,789; this figure exceeds average annual reported hospitalized mumps disease costs of $61,525 by a factor of 4.7. Mumps attack rates would have to reach at least 15.0 per 100,000 per year before savings would equal recruit vaccination costs. Failure to show that a vaccine program would be cost-saving may be due to limitations in identifying cases or to the requirement that all recruits be immunized regardless of prior immune status. It is likely that a program to immunize susceptible individuals alone would show benefit. (*Am J Public Health* 1989; 79:471–474.)

Introduction

Mumps has been perceived as an acute communicable disease of childhood for at least 20 years, but in earlier times it was viewed primarily as an illness that affected armies during mobilizations. Mumps was one of the most important causes of days lost from active duty by the US Army in France during World War I, with an average annual hospitalization rate of 55.8 per 1,000 average strength.[1] In World War II and afterward, however, the US military experienced much lower mumps rates and only isolated epidemics. This decrease between the world wars was attributed to increasing urbanization with resulting increased transmission during childhood.[2]

The introduction of mumps vaccine in December 1967 resulted in a dramatic decline in reported mumps attack rates among the US population. Centers for Disease Control (CDC) figures show a 98 per cent drop from 1968 levels, to an all-time low of 1.2 cases per 100,000 population in 1985.[3] In 1986 and 1987, however, there were reports of mumps outbreaks in secondary schools and colleges.[4,5] US rates rose to 3.2 per 100,000 in 1986, with 7,790 cases reported, while the provisional total for 1987 is even greater at 12,200 cases.[5] These increases have resulted in calls for vaccination programs directed at young adults who have not been immunized previously.[4,6]

The US Army, whose soldiers are largely in their early 20s, experienced a similar mumps outbreak in a unit stationed in the Republic of Korea during the spring of 1986. This outbreak consisted of 29 known symptomatic cases over a four-month period, with a crude attack rate of roughly 4 per 1,000 soldiers. (CPT Robert Pero, MD, unpublished data.) In response to this outbreak and the perception that a new immunization program for recruits might be indicated, this study was undertaken to examine the recent history of mumps in the US Army, and to consider whether an immunization program would be justified on cost grounds.

Methods

The US Army Medical Department maintains a worldwide, computerized Individual Patient Data System (IPDS) hospital admission data base for administrative purposes. Each IPDS hospital discharge record contains basic demographic, geographic, length-of-stay, and disposition information, including convalescent time, as well as up to eight diagnoses and eight procedures. These are coded according to the International Classification of Diseases, 9th Revision (ICD-9).

All discharge records with a mumps-related ICD-9 code for the years 1980–86 were retrieved. This period was chosen because it was felt that policy should be based on the Army's most recent experience with mumps illness for which accurate data are available. Because the question of interest is mumps disease impact on active duty soldiers, the 215 total cases identified were limited to a subset of 152 active duty Army, excluding dependents, retirees, and other service members. Denominator data were obtained from the Defense Almanac, published annually by the Department of Defense.

While mumps in the civilian population may be underreported by as much as a factor of 10,[7] military symptomatic cases have a much higher reported hospitalization rate, especially among lower ranking soldiers who live in barracks. These individuals are admitted to respiratory disease wards largely as a transmission control measure, e.g., virtually every overtly symptomatic case in the Korean outbreak was hospitalized. Hospitalized cases were therefore used as a proxy for all cases, with the understanding that some underestimation remains.

Outpatient costs associated with mumps hospitalization were calculated assuming two clinic visits per case (admitting and follow-up) plus one additional follow-up visit for each complication reported during hospitalization. Clinic and daily hospitalization costs for the most recent (1986) fiscal year were obtained from the Army Health Services Command (HSC) which calculates average costs across all Army hospitals within the 50 states. These costs are $322.82 and $67.83, respectively, and include all associated plant, salary, supply, and inpatient medication expenses. Overseas costs in US Army facilities are similar, due to use of similar staffing and military supplies.

As the analysis used is based on a steady state model in the present, the 1986 costs were applied to all cases, ignoring inflation during the period. Intangible costs for pain and suffering were excluded. Average annual costs were computed using the mean annual number of hospital days and convalescent days (to estimate total lost pay costs) for the entire seven years of data, and these average costs were

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compared with average annual estimated vaccine program costs.

The Army currently administers measles and rubella vaccine (MR) to all enlisted recruits. Therefore the cost to initiate a mumps vaccination program would be little more than the Army's additional cost per dose for measles, mumps and rubella vaccine (MMR) over MR. In mid-1987, this cost was $1.542, based on the standard 10-dose vials. The estimation of mumps vaccine administration costs was derived by applying this incremental cost to the 180,000 average annual number of enlisted accessions to basic training (MAJ John Brundage, MD, MPH, personal communication), a number which includes Army Reserve and National Guard enlistees.

An 85 per cent overall vaccine effectiveness rate was assumed in calculating disease costs that would accrue in spite of a vaccination program. Recent estimates of clinical mumps vaccine efficacy range from 75–91 per cent.[5,8,9] Officer accessions, however, as well as all current active duty would not be included in the proposed program, so overall effectiveness in this population would tend to be less than the maximal efficacy. (In the earliest years of such a program, effectiveness might be lower still, since only new accessions would be covered.)

Results

Table 1 summarizes the active duty experience of mumps hospitalizations over the seven-year period. While grouped age-specific rates are included for comparison, the Army's age distribution is very skewed to younger ages in the groupings above age 20, causing some of the apparent elevation of these rates. The direct age-adjusted Army rates are below reported US rates for the same period. Prior to 1986, rates had been trending downward for both the Army and the US, but the Korean mumps outbreak caused the Army rates to increase five-fold that year as compared to 1985. Through December of 1987, the Army recorded 33 active duty cases, making the provisional 1987 rate 4.2 per 100,000. This is a decline from 1986 rates, yet is still twice the 1980–85 average of 2.14.

TABLE 1—Annual Incidence of Mumps Hospitalizations per 100,000 Active Duty US Army Soldiers for the Years 1980–86 Compared with Overall US Mumps Rates as Reported by the CDC*

Incidence	1980	1981	1982	1983	1984	1985	1986
Army cases	30	16	10	18	16	10	52
Army rate	3.85	2.05	1.28	2.31	2.05	1.28	6.65
Age-adjusted Army rate†	2.2	0.9	0.5	0.9	1.0	0.4	2.7
Army rates by age group							
16–20	7.5	3.2	3.1	4.3	3.1	2.6	14.1
21–25	2.2	2.2	1.1	2.3	2.7	1.9	7.3
26–30	3.8	0.8	0.7	2.0	0.7	0.6	3.8
31+	2.8	1.7	0.5	1.0	1.6	0.0	2.5
US rate	3.9	2.2	2.5	1.6	1.3	1.2	3.2
US rates by age group							
15–19	2.6	2.0	3.8	1.8	2.0	2.4	9.0
20–24	0.6	0.6	0.6	0.4	0.4	0.5	0.9
25–29	0.5	0.5	0.5	0.3	0.3	0.4	0.4
30+	0.2	0.1	0.1	0.2	0.2	0.1	0.2

*US rates for 1980–83 derived from MMWR's Annual Summary, 1980 through 1983, published by CDC; other US rates from reference 5. DOD and CDC use different age groupings in compiling their published data.

†Direct adjustment to resident US population distribution per Statistical Abstract of the United States, published annually by US Bureau of the Census.

Age, race, sex, and length of service-specific rates for the entire period are presented in Table 2. There is a clearly decreasing trend in attack rates with respect to age and length of service. While rates for Blacks and Whites appear similar, those for other minorities are significantly higher. In 1985, the Army began classifying Hispanics as a distinct race/ethnic category in IPDS data. Race/ethnic rates were recalculated for 1986 alone, with the Hispanic rate, based on five cases, being more than twice that of Whites. No overall difference with respect to sex was found.

Clearly, the sharp decline in rates with increasing length of service is confounded by the fact that those with longer service records are usually older. However, mumps is not occurring primarily in basic training camps. Only 15 (9.9 per cent) of the cases occurred in soldiers within their first 90 days of service.

It was considered that because of much higher rates 1986 might have been different from the 1980–85 period, so the analysis was repeated for these years separately. Except for the higher rates, the mumps disease patterns in 1986 were consistent with the prior years, and it is therefore appropriate to use all seven years of data for the overall cost-benefit calculations.

Complication rates are important in estimating permanent sequelae and total disease costs. Unfortunately, the ICD-9 does not list all possible mumps complications in its coding system (e.g., mastitis, reported in 15 per cent of females[10]) so it cannot be determined if these occurred. Seventy-nine per cent of the cases had no reported complications. Among mumps-related complications, orchitis was seen most frequently (26 cases, 19 per cent of males), followed by meningitis (two cases) and encephalitis, nephritis, and testicular atrophy (one each). These complications were apparently mild, and all 152 patients returned to active duty without disability.

The median number of sick days (inpatient plus conva-

TABLE 2—Summarized US Army Age-, Race-, Sex-, and Length of Service-Specific Hospitalized Mumps Cases and Rates per 100,000 Active Duty for 1980–86

	Cases	Rate
Overall total	152	2.78
Age group* (years)		
16–20	63	5.54
21–25	52	2.81
26–30	19	1.77
31–35	9	1.34
36–40	6	1.38
41–45	2	1.14
46–50	1	1.72
51	0	0.00
Race		
Black	35	2.29
White	95	2.66
Other†	22	5.56
Sex		
Male	139	2.84
Female	13	2.16
Length of Service (years)*		
<1	57	7.05
1–3	56	2.99
4–6	17	2.19
7–9	7	1.19
10+	15	1.16

*$p < 0.0001$ by χ^2 test for trend.

†Includes Hispanic cases, which were not reported separately prior to 1985. Hispanic rate for 1986 alone = 18.3 per 100,000.

lescent time) was seven, and the mean 9.6. Inpatient care amounted to 73 per cent of total sick time, with the remainder spent on convalescent leave. The range in total sick time was 1 to 267 days, with 95 per cent of the patients having 19 or fewer days. The most extreme outlier suffered complications due to gastric surgery. Another outlier at 49 days suffered complications of labor and delivery. Since it was not known whether these complications were not in some way related to the individuals' mumps illnesses, they were retained in the cost calculations with the understanding that any resulting bias would tend to favor a decision to immunize.

Estimated annual disease costs were calculated based on the mean attack rate for the 1980–86 period. There were a total of 1,057 inpatient hospital days recorded, an average of 151 per year. Multiplying by the specified daily hospital cost yields an average annual hospitalization cost of $48,764. The estimated total number of associated clinic visits was 207, or 30 per year. Multiplying by the specified clinic visit cost yields an average annual clinic cost of $2,006. Lost pay costs based on total reported sick days and pay scales in effect on January 1, 1987 were $75,412, an annual average of $10,773. The sum of these three average annual costs is the total estimated disease cost for an average year at 1986 cost levels, $61,525.

Estimated vaccine program costs (annual number of enlisted accessions multiplied by the $1,542) yielded a figure of $277,560. Applying the estimated 85 per cent vaccine effectiveness rate to the average number of reported cases left an expected three nonpreventable cases. Costs for these cases were calculated proportionally to be $9,229. Thus, the total estimated annual expenses under a vaccination program would be $286,789. This exceeds the direct and indirect costs for hospitalized cases by a factor of 4.7.

Table 3 compares disease costs with vaccination program costs at various attack rates assuming a steady state. The costs for 1986 were calculated from that year's data alone by the same procedure as the overall comparison described above. It is worth noting, however, that the extrapolated estimate of total 1986 disease costs, using the overall average cost per case, would be within $2,500 (1.6 per cent) of the actual cost. This agreement tends to support the extrapolation to the break-even number of cases and corresponding attack rate, which were solved for algebraically under the assumption that all disease costs would rise proportionately with increasing rates.* Reported mumps attack rates would have to exceed at least 15.0 per 100,000 annually before the cost ratio would favor vaccination.

TABLE 3—Comparison of Hospitalized Disease vs Vaccination Program Costs at Various Mumps Attack Rates

	Average 1980–86	1986	Break-even*
Attack rate (per 10^5 per year)	2.78	6.65	15.0
Number of cases	22	52	117
Mumps disease cost			
Medical costs	$50,752	$126,240	$269,350
Nonmedical costs	10,773	16,977	57,200
Total	$61,525	$143,117	$326,550
Vaccination program costs			
Vaccine cost	$277,560	$277,560	$277,560
Non-preventable case costs	9,229	21,468	48,990
Total	$286,789	$299,028	$326,550
Vaccine:mumps cost ratio	4.7	2.1	1.0

*Projected costs and rate at break-even point are rounded estimates.

Discussion

The use of hospitalization records to estimate rates and compute costs is not without pitfalls, but some of these may be minimized by the unique aspects of the military population under study. Because of the much higher hospitalization rates for mumps in the military, these rates are felt to be an adequate proxy for reported rates. However, there are still limitations in identifying cases. Asymptomatic cases are almost never reported in either the civilian or the military sector, but from a cost-benefit standpoint, asymptomatic cases do not contribute any costs, and their exclusion is less important. Mild cases, which are unhospitalized and undocumented, would incur medical and nonmedical costs, but at a lower rate.

The racial differences noted may be of some interest, especially the apparently high non-Black minority rate. Previous studies have tended to show that race has no effect, although Blacks may have higher rates at an earlier age.[11] Our data are insufficient to determine if these higher rates are limited to Hispanics alone. A similar data analysis showed that in 1985 Hispanics in the Army had roughly six times the rate of varicella illness, relative to Whites (MAJ Patrick Kelley, MD, MPH, unpublished data). Many of these cases came from Puerto Rico where varicella transmission is low. It is speculated that either a similar difference in prior exposure or a difference in prior immunization is occurring with Hispanic mumps cases.

As any mumps-related diagnosis caused inclusion in the data set, it was impossible to determine if a complicated case acquired the infection before or after admission, and thus whether the reported complications were directly related to the mumps infection. The decision to include the outlying cases in the cost calculations increased the estimated disease costs by roughly 20 per cent. On the other hand, the fact that the IPDS data did not differentiate intensive care unit (ICU) bed days from regular bed days might lead costs to be underestimated (the Army's medical ICU bed cost is nearly three times higher than regular bed cost), although medical ICU usage in mumps illness is probably limited to the rare severe meningitis or encephalitis case.

Including Army Reserve and National Guard trainees in the proposed immunization program increases program costs, while most disease prevention savings would occur outside the military health care system where it is not captured in the calculations. These recruits train side-by-side with active duty soldiers and are required to receive the same vaccinations, but return to civilian life within 120 to 180 days. About one-third of the basic trainees are reservists or guardsmen during a given year. Assuming their disease rates are the same as the active duty population, the total mumps disease cost might be one-third higher. This is still not enough for the program to be cost-beneficial.

The conclusion that the proposed vaccination program might not save money was not expected. There is no doubt

*The equation solved is (cost per case) x = (vaccine cost) + (proportional cost for non-preventable cases), where x is the number of mumps cases at which vaccination program costs equal disease costs, and the proportional cost for non-preventable cases is 0.15 (cost per case) x. This is a linear extrapolation based on the average cost per case.

that vaccination programs are cost beneficial when applied to completely unvaccinated populations.[12,13] What is certainly the case here, however, is that a large proportion of the military recruit population is already immune, due to prior vaccination or illness. Surveys of US military recruits taken in the 1950s and 1960s, when post induction mumps rates were higher, found antibodies to be present in 54–76 per cent.[14,15] Cochi, *et al.* using data from the United States Immunization Survey, estimate vaccine coverage in 15–19 year-olds at 55 per cent and rising, as of 1985.[5] This immunity has two major impacts on the results. First, it prevents attack rates from approaching those seen in pediatric or otherwise virgin populations, on which most previous cost-benefit analyses were based. Second, it means that a large amount of the vaccine given in a blanket immunization program would be wasted, raising vaccination costs unnecessarily.

The Army does not generally use previous civilian immunization records to exempt basic trainees from the usual schedule of vaccines, so the obvious solution of simply vaccinating those without proof of prior immunization could not be immediately implemented. Relatively few recruits now arrive at basic training camps with valid immunization records in hand, and the administrative costs of determining each individual's immune status may exceed the $1.54 vaccine cost. As preinduction immunization rates for mumps and other diseases rise, however, a policy of selectively immunizing recruits based on review of records would likely become cost-beneficial.

Excluding intangible costs for pain and suffering, as well as neglecting later life costs of sequelae, may be questioned in this study because of the failure to show any strict monetary benefit from a vaccine program. In fact, if the cost ratio had been closer to 1.0 we would have argued in favor of a vaccination program on the basis of such benefits. But the actual complications seen here and those generally reported[4,10] are usually self-limited, and serious sequelae are generally considered rare,[12,13,16] so it is less likely that this analysis has overlooked a large potential cost that would invalidate the conclusions. Reports of sterility, perhaps the most feared sequela in postpubertal males, are very sparse,[16] and it is therefore discounted.

The usual explanation for the recent rise in mumps incidence among secondary school and-college age populations is that there is a nonimmune cohort in the population that missed both mumps immunization and illness in the 1970s, when rates were falling and before states passed school vaccination laws which included mumps.[5,9] If this is the case, the frequency of mumps outbreaks should decrease after a few years, as this cohort passes into adulthood. In the interim, monitoring of Army mumps rates for any increase is indicated, and the 16 states that currently do not require mumps immunization for grade school enrollment[17] should be encouraged to enact such laws. While this study's perspective is necessarily narrow, in that only current disease costs are considered, it appears that Army mumps rates would have to more than double from 1986 levels before a blanket immunization program would be cost-beneficial. A current serological survey of the recruit population is planned to better address potential changes in requirements for mumps and other vaccines.

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The Role of Secondary Vaccine Failures in Measles Outbreaks

Richard G. Mathias, MD, William G. Meekison, MD, Teri A. Arcand, and Martin T. Schechter, MD, PhD

Abstract: An outbreak of measles in 1985–86 in a community where measles vaccine trials had been carried out from 1974–76 allowed the assessment of the role of secondary vaccine failures in previously immunized children. A total of 188 children from the vaccine trial were followed. Of these, 175 seroconverted initially while 13 (6 per cent) required re-immunization (primary failure). A total of 13 cases of measles, eight of which were laboratory and or physician-confirmed, were reported in this cohort. Of these, nine cases occurred in the 175 subjects who had hemagglutination inhibition test (HI) and neutralizing antibody responses following the initial immunization. These nine cases represent secondary vaccine failures. An additional four cases occurred in the 13 subjects with primary vaccine failure. We conclude that secondary vaccine failures occur and that while primary failures account for most cases, secondary vaccine failures contribute to the occurrence of measles cases in an epidemic. A booster dose of measles vaccine may be necessary to reduce susceptibility to a sufficiently low level to allow the goal of measles elimination to be achieved. (*Am J Public Health* 1989; 79:475–478.)

Introduction

Measles virus has probably infected mankind since antiquity. Only recently has the introduction of a vaccine allowed the control of measles to be considered a goal of public health.[1–3] Panum[4] showed that immunity to wild measles virus was maintained even many years after infection. However, immunity following vaccine has been more problematic, with outbreaks of measles described in immunized populations virtually since the vaccine was introduced and widely used.[5–16] In all of these reports, failure of vaccine delivery was thought to be the major factor in the outbreak rather than failure of the vaccine itself. In those cases who had been immunized and developed measles, the reason was felt to be lack of seroconversion with initial immunization, primary vaccine failure, rather than loss of protection after seroconversion had occurred, secondary vaccine failure. The attack rates in the immunized of about 5 per cent were compatible with primary vaccine failure rates. Those individuals who do not seroconvert to measles vaccine and hence are possible primary vaccine failures may be detected by serology and given further vaccine if desired. Secondary vaccine failures who have seroconverted but are not protected cannot be so identified. Gustafson[17] has recently described an outbreak in Texas in a well-immunized population where primary vaccine failure was thought to be a more important factor than vaccine delivery.

We had the opportunity to study the incidence of measles in a population of children who had participated in measles vaccine trials and who had antibody studies carried out following immunization. We report here on the relative importance of primary and secondary vaccine failures on the incidence of self-reported measles in a well-immunized population in the 10 years following immunization. As present and future control of measles is highly dependent on the induction of immunity and the maintenance of protection to prevent measles,[18] these findings may indicate the possible success or failure of measles elimination programs.

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The Setting

Boundary Health Unit (BHU), a public health unit funded by the Ministry of Health of the Province of British Columbia, serves a suburban and rural community close to the major metropolitan center of Vancouver. In 1985, the year of the start of the measles outbreak, there were 271,141 residents in the 480 square kilometer area administered by the unit. Of these residents, 8.1 per cent were under age 5, 16.7 per cent were aged 5–14 years, and 31.7 per cent were aged 25–44 years. The mean family income was $29,008. The mean family size was 3.0 individuals. There are 124 public and private schools with an enrollment of 34,218 in kindergarten to grade 7, and 21,178 in grades 8–12.

Vaccine trials were carried out from January 1974 to February 1976. There were approximately 6,000 one-year-olds in 1974–75 eligible for measles immunization at one year of age. Of the number potentially eligible, 382 volunteered for the study and were enrolled. Connaught Laboratories Limited of Willowdale, Ontario conducted the measles vaccine trials to compare vaccine produced by them to vaccine from Merck, Sharp, and Dohme, Ltd. Both vaccines contained a live further attenuated measles strain and both were given as a single product. Children were immunized at approximately one year of age after an initial serum specimen was drawn. A second sample was drawn one month later. Repeat samples were done in the immunized children who could be located at three and five years. All serum samples obtained in the vaccine trial were tested for hemagglutination inhibition and for neutralizing antibodies by Connaught.

While British Columbia has experienced a decline in the numbers of measles cases reported, outbreaks have continued to take place. A measles epidemic occurred in the province from 1984 to 1986 involving 7,479 reported cases for the 12 months from September 1985 to August 1986. Of these, a total of 744 cases were reported from the BHU (354 males and 390 females); 60 (8.5 per cent) were confirmed by the Virology Section of the Provincial Laboratory, BC Ministry of Health. The peak of the epidemic occurred from the end of January 1986 to the middle of March 1986 with cases occurring until the second week of June 1986. The mean age of reported cases was 12 ± 5.6 years. The largest proportion of cases (42.7 per cent) occurred in the age group 10–14 years.

Methods

During May–June 1986, all children who had taken part in the 1974–76 vaccine trials were identified from the trial

center's records. An attempt was made to contact them at their last known address using school lists to locate additional children. The child and his/her parents were contacted to obtain consent to participate in the follow-up study. For those who consented, a blood sample was obtained from the child, and a questionnaire was administered by study personnel to the child and a parent. The questionnaire sought information about the immunization history and possible cases of measles in the child and in family members. The children in the vaccine trial who were located and enrolled in the follow-up study will be referred to as the cohort.

The measles hemagglutination (HI) antibody levels and the measles neutralization titers were both determined by methods described by Gershon and Krugman.[14] The measles IgM (immunoglobulin M) levels were determined by a method adapted by Connaught Laboratories Ltd., and are on file.

Antibody levels were analyzed using geometric mean titer (GMT) as doubling concentrations had been used for all specimens in the vaccine trial. The samples were run in different years. Although the methods for HI and NA were similar over the time period, variations in materials and instruments makes direct comparison of levels in different years inadvisable. This is particularly true of the initial and 1986 samples. We have thus looked at samples within years and did not analyze titers serially.

Primary vaccine failure is defined for this paper as failure to have neutralizing antibodies following immunization, whether the first or subsequent immunization. Secondary vaccine failure is defined as a lack of protection in an individual who had documented neutralizing antibody titer ≥ 1:4 post-immunization. Seroconversion is defined as a four-fold or greater rise in hemagglutination titer, or a detectable titer in an individual who had no previous detectable antibodies.

Results

A total of 382 infants had been enrolled in the vaccine trials. Due to an unbalanced design, 265 received vaccine produced by Connaught, while 117 received measles vaccine produced by Merck. A total of 225 of the 382 families were contacted. Thirty seven gave information but no sample was obtained. Thus for the follow-up study, a total of 188 subjects (49.2 per cent of the vaccine trial participants) were located and had data and blood samples collected (cohort), while the remaining 194 (51.8 per cent) were not located or did not enter the follow-up study (lost). Of the cohort, 129 and 59 members had received Connaught and Merck vaccines respectively, the same ratio as in the original trial. As seen in Table 1, the cohort and lost groups were similar.

The proportion of the cohort and lost groups subsequently followed in the vaccine trials was 26 per cent of the cohort group and 8 per cent of the lost group recalled at a mean of 58 months for both groups. The titers for both groups at each follow-up period were similar except for higher HI and NA titers in the cohort group at 30.4 and 31.5 months respectively. By 58 months, the antibody titers in both groups were similar.

The mean age at immunization was 1.04 years for the cohort; 23 cohort members had more than one immunization; 13 were failures in the cohort group and 12 individuals were reimmunized as part of outbreak control measures. Seroconversion rates were similar for both vaccines.

Thirteen children in the cohort group (6.9 per cent) were reported to have had measles subsequently. One of these cases was laboratory confirmed. Neutralizing antibody had been demonstrated in this individual post initial immunization. Eight cases of measles in the cohort were confirmed by a physician. Among the 13 cases of measles, five had received Merck vaccine, eight Connaught. Of these 13 cases, two had measles at age 2, four at ages 6–8 (the 1980 outbreak), and seven at ages 10–12 (the 1985–86 outbreak).

As seen in Table 2, of the 13 subjects in the cohort who did not seroconvert initially to NA, 10 underwent repeat immunization with antibody testing and all 10 seroconverted. One of these 10 was reported to have had measles. There were an additional two infants who had neither a second immunization nor subsequent antibody titers (primary vaccine failures) and both of these children were reported to subsequently have had measles. The remaining child with measles received a repeat immunization but seroconversion was not documented; the latter child was also reported to have had measles. Overall, four of 13 (31 per cent) of those who were initial vaccine failures subsequently had measles. These four cases were all physician-confirmed and three occurred in 1985–86. In those whose follow-up antibody levels were considered protective, nine of 175 (5 per cent) subsequently had measles. These nine cases were secondary vaccine failures. Four of the nine cases were physician-confirmed and four were diagnosed in the 1985–86 outbreak. The attack rates (and 95% confidence intervals) in those who converted on initial vaccination and those who did not were 31 per cent (9, 61) and 5 per cent (2, 10), respectively.

In the cohort, the initial laboratory titers differed between those with and without subsequent measles (Table 3) in that both HI and NA titers were lower in the group who subsequently developed measles. This was observed both for those who had ever had measles and for those who had measles in the 1985–86 epidemic.

The results of the samples taken in 1986 indicated that the mean HI titers in the cohort members who reported measles was higher than in those who did not (geometric mean titer 30 vs 19). This was even more pronounced if only

TABLE 1—Comparison of 188 Children Followed from Original Vaccine Trial (cohort) and 194 Children Lost to Follow-up (lost)

Parameter	Cohort (%)	Lost (%)
Number	188	194
Mean Age at Immunization	1.04 ± 0.10	1.06 ± 0.14
Seroconversion: HI	177 (94.1)	185 (95.4)
Seroconversion: NA	175 (93.1)	185 (95.4)
Total Seroconversion	185 (98.4)	193 (99.5)

TABLE 2—Reported Measles Cases in the 188 Cohort Children Stratified by NA Response to Measles Vaccine

Initial response	Total	Number Reporting Measles	Per Cent Attack Rate (95% CI)
Initial non-conversion	13	4	31
Repeat immunization with conversion	10	1	(9, 61)
No repeat immunization	2	2	
Repeat immunization conversion unknown	1	1	
Initial seroconversion*	175	9	5
			(2, 10)

*Seroconversion defined as NA greater than 1:4

TABLE 3—Geometric Mean Titers for the 188 Cohort Children Stratified by Reporting of Subsequent Measles

Measles History	Geometric Mean Titers*		
	Initial HI	Final HI	Initial NA
Ever had measles			
yes	14.4	30.2	8.9
no	36.3	18.6	19.5
Measles in 1985–86			
yes	8.8	40.3	7.2
no	35.8	18.7	19.2

*Initial refers to the sample following immunization in the vaccine trial; final refers to the 1986 follow-up sample.

measles cases with dates of onset in 1985–86 were used (40 vs 19).

The measles specific IgM results were negative for all 188 members of the cohort. Specifically, none of the seven who reported measles in the 1985–86 outbreak were IgM positive. The samples for the IgM determinations were taken 13 and seven months after reported disease in two individuals, four months post infection in three subjects, two months in one, and two weeks in the other. A positive result may have been expected in five of these cases.

Discussion

Although these children were originally enrolled in a vaccine trial, their subsequent experience with measles does not appear to have been due to the differences in vaccines being tested. The control product from Merck has been widely used in Canada and the US at that time and since. Although the Connaught vaccine was not marketed, the proportion of cases who seroconverted and who subsequently developed measles was not significantly different from the Merck immunized children. The similarities in the located and not located groups within the vaccine trial population indicate that a bias in the follow-up is not likely to account for the experience of the cohort with respect to measles.

All of these children had neutralizing antibody detected postimmunization although only 95 per cent seroconverted with the first immunization as might be expected.[18,20] All of these children would conventionally be considered protected including those who required re-immunization.[21] In those in whom seroconversion had been documented, any subsequent measles must be considered a secondary vaccine failure as defined by Frank, *et al*.[22] Although documented by Cherry,[5] Frank, *et al*,[22] did not consider secondary failure a major contributor to the measles elimination program failures.

The measles experience of this cohort was considered over the life experience of the child. With the difficulty in making the clinical diagnosis of measles, this may result in the overestimation of the number of cases. Nevertheless, the reported cases in the cohort took place during reported measles outbreaks in British Columbia. During the recent outbreak there was one laboratory-confirmed and a number of physician-confirmed cases reported. Following the measles, HI antibodies demonstrated a rise over time in those reporting measles compared to those not reporting measles. All of these observations would tend to validate the cases reported. We conclude that the reported cases in the cohort represent measles in children who had seroconverted and hence are secondary vaccine failures.

Within the cohort, there were predictors of subsequent measles. Children who required re-immunization because of an initial vaccine failure were six times as likely to have measles as those who seroconverted with a single dose. Even after seroconversion, the mean titers of seroconverters who reported measles in the 1985–86 outbreak were less than those who did not report measles (HI 35.8 vs 8.8; NA 19.2 vs 7.2). These data indicate that the initial immunization response predicts subsequent risk of measles and that the risk depends on both seroconversion and on the titer of antibody produced at approximately one month following immunization. These low titers may become undetectable more quickly than higher titers although other investigators have not found that time from immunization was a factor in measles outbreaks.[6,23–25]

In our cohort, none of the five recent cases was found to be IgM positive. Nagy, *et al*.[23] found IgM responses in 77.4 per cent of those who had been vaccinated. They concluded that those who had an IgM response represented a primary vaccine failure, whereas those with HI increases or a high titer were secondary vaccine failures. Gustafson, *et al*,[17] found cases only in the group seronegative early in the epidemic, with all appropriately tested cases IgM positive. They concluded that, as no cases occurred in the seropositive group, the outbreak they investigated was due to primary vaccine failures. The situation in our outbreak was markedly different as 184 of 188 individuals in the cohort were known to be seropositive following immunization. Although it is possible that all of the cases of measles occurred in those who had lost antibody, as explained by the lower initial titers in the cases,[24,25] we could not confirm this. The lack of IgM response in our group is in contrast to the data of Gustafson[17] and Nagy,[23] although the very small numbers in our group make this lack of an IgM response imprecise. The pre-existing antibody titer results on our cohort confirm that secondary vaccine failure was important in our study population.

The ultimate control of measles has been predicted with models developed by Anderson and May[26] and Levy[16]. The role of secondary vaccine failure has not been considered in these models. Neither was it considered important by Frank, *et al*.[22] We feel that the documentation of an attack rate approximating 6 per cent over the 10–12 years post-immunization in a cohort of individuals with known primary vaccine seroconversion must prompt further study into the role of secondary vaccine failure in measles outbreaks. Nkowane, *et al*,[6] found that at least 48 per cent of their cases occurred in adequately immunized individuals. Although the outbreak they investigated lasted only four generations, they concluded that vaccine failures played a role in transmission.

If a significant role is established for secondary vaccine failure, this must be taken into account in the predictions of measles control programs. The assumption that measles immunity as induced by vaccine is as high as the seroconversion rate appears to be an overestimate of the true situation. Routine booster immunization of all children rather than for select groups,[24] in order to reduce the numbers of primary vaccine failures and to boost the titers in all children, may be necessary to eliminate measles. It is clear that reliance upon measles antibody studies, even when neutralizing antibody is measured, only gives an estimate of the immune status of the individual. The clarification of the constituents of immunity to measles may allow the develop-

ment of vaccine which will more closely stimulate the same immune response that is induced by wild measles infection. Until this is done, it appears that we will continue to see measles epidemics in well-immunized populations and that these epidemics will be due to both primary and secondary vaccine failures.

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New Fellowship to Honor Creativity in Managed Health Care

An annual fellowship grant has been established by the Caldwell B. Esselstyn Foundation to honor an individual who demonstrates exceptional creativity and scholarship in the field of managed health care. To be awarded for the first time in July 1989, the fellowship—which could total as much as $35,000—will provide seed money for development of a new project or program to benefit staff or group model HMOs. The foundation is named for a pioneer in the prepaid group practice movement.

To obtain an application for the fellowship, write or call the Caldwell B. Esselstyn Foundation, 1201 Troy-Schenectady Road, Latham, NY 12110. Tel: (518) 783-1864. Completed applications must be submitted by May 15, 1989.

Illness Associated with Child Day Care: A Study of Incidence and Cost

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Abstract: We studied 843 children under 36 months of age enrolled in a prepaid health plan from September 1985 through March 1986, to identify characteristics of day care which might be risk factors for infection and to describe the resulting economic costs. Children cared for in their own home had a mean of 2.03 infections diagnosed during the study period. Adjusted rates of excess infection (95 per cent CI) for children cared for in other settings were: −.09 (−.73, .54) in relatives' homes; .10 (−.51, .71) in day care homes; .79 (.13, 1.45) in day care centers; .60 (−.24, 1.46) in mother's day out programs; and .66 (−.01, 1.34) in multiple settings. Children in day care centers were 4.5 times more likely to be hospitalized than those in other settings (95 per cent CI = 1.55, 13.00), primarily due to an increased rate of tympanostomy tube placement (relative risk 3.79, 95 per cent CI = 1.04, 13.36). The strongest predictor of illness risk was the number of other children in the room. The mean monthly cost of medical care was $32.94 for children in the highest risk settings compared with $19.78 for those in other settings. Illness in a child in our study accounted for 40 per cent of parental absenteeism from work; the mean number of days lost per month was 0.52 for parents of children in day care centers compared with 0.37 for those of children in other forms of full time care outside the home. (*Am J Public Health* 1989; 79:479–484.)

Introduction

Day care for preschool children of working parents has become an integral part of American life. More than half of all mothers with preschool children are in the labor force and more than 60 per cent of their children receive day care outside the home.[1] Many studies show an increased risk of infection for children and their families using at least one form of care—the institutional day care "center".[2–10] A lower risk of illness has been attributed to children cared for by non-relatives in private households known as day care "homes", but few data are available from these settings which are typically small, often unlicensed, and difficult to identify.[5–9]

Reliable measures of the risks and consequences of infections attributable to various forms of child care are needed. Most reports have focused on only a single pathogen or disease, studied day care centers exclusively, or used case-control methods to estimate relative risk of illness rather than prospective surveillance to generate incidence data. With some exceptions,[11–14] most studies have not evaluated specific characteristics of child care which might promote the spread of infection. Finally, despite the fact that child care has become an economic necessity for many families, the economic consequences of varying illness rates among children in different settings have not been examined.

During the winter of 1985–86, we studied a cohort of preschool children in our practice who were enrolled in a prepaid health care plan to compare the incidence of illness among children cared for in different settings, to study the economic consequences of child care-related illness, including increased utilization of health care services and parental absenteeism from work, and to identify characteristics of child care which may be important determinants of illness incidence and health care costs in preschool children.

Methods

The children in our study received prepaid medical care from the Health First Medical Group (HFMG), a 28-physician multi-specialty group having three offices in middle class areas of Memphis, Tennessee. The children studied were all insured under a family health care plan offered through a parent's employer which stipulated that all medical care, in order to be covered, had to be provided or authorized by an HFMG physician. Parents, employers, or both paid a single monthly premium, in exchange for which the costs of all visits to an HFMG office, including laboratory and x-ray studies, as well as referrals to specialists and hospitalizations were covered at 100 per cent with no deductible. Authorized emergency room visits were also covered but incurred a $25 copayment. Prescriptions written by HFMG physicians could be filled immediately at in-house pharmacies for a fee of $3 each. These features were designed to provide strong incentives for patients to obtain all of their medical care from the HFMG.

An HFMG office is located within 20 minutes driving distance of all areas of Memphis and its suburbs. To facilitate access to medical services, HFMG offices are open weekdays from 8:30 am until 9:00 pm, and weekends and holidays from 1:00 to 5:00 pm. A 24-hour telephone consultation service is also available.

Study Design and Data Collection

During the seven-month period September 1985 through March 1986, health care utilization data for all children under 36 months of age enrolled in the prepaid plan were collected prospectively without their parents' knowledge that a study was in progress. Information regarding illness diagnosis and laboratory and x-ray studies (obtained if necessary for patient care) were recorded on computerized encounter forms by the examining physician at each HFMG office visit. For this study, common infectious illnesses were classified as: 1) upper respiratory infection without otitis media; 2) lower respiratory infection; 3) acute otitis media; 4) gastroenteritis; and 5) skin infection. We distinguished follow-up visits for the same illness from separate illnesses.

Medical charges were abstracted from the encounter forms; these were charges which would have been billed directly to other HFMG patients not members of the prepaid health care plan. Diagnoses and charges resulting from authorized emergency room visits and hospitalizations were obtained subsequently from hospital and insurance records. The average retail cost per visit of drugs prescribed for children during the study period was obtained from pharmacy

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records. Immediately following the study period, a letter explaining the project and a questionnaire were mailed to parents of all subjects. The questionnaire requested information on child care arrangements for the index child, including household characteristics and other variables which might affect the child's risk of illness (Appendix A). We also asked the number of days and the reasons for which a parent had been absent from work during the preceding three months, the cost of day care for the child, the number of visits the child had made to non-HFMG physicians or emergency rooms during the preceding month and the number of times the child had been ill without seeing a physician during this month. Two weeks after the questionnaires were mailed, a computer generated random sample of 15 per cent of parents who had not responded were interviewed by telephone. This survey was performed by trained interviewers who explained the purpose of the study and administered the questionnaire using a standardized, closed-ended format.

We excluded from the study only those children diagnosed by one of us as having a chronic illness which might lead to increased susceptibility to infection (including asthma, bronchopulmonary dysplasia, congenital heart disease, cystic fibrosis, and Down's syndrome).

Tympanostomy tubes were placed by a single otolaryngologist after referral by one of us for indications of either recurrent episodes of acute otitis media despite daily prophylaxis with antibiotics, or middle ear effusions persisting despite three months of medical treatment. Patients who had received tympanostomy tubes within eight months before onset of the study were excluded from the tabulation of otitis media incidence; patients who received them during the study were excluded from this tabulation for the remainder of the study.

Statistical Methods

Bivariate analysis of illness incidence according to type of child care was performed using the F test for analysis of variance. Rates of excess illness in children cared for outside of their own homes were adjusted using linear regression analysis for the effect of host and family environment variables which differed among children in different forms of care.

Hierarchical multiple linear regression analysis was used to evaluate potential risk factors for diagnosed illness in children who were cared for outside of their homes by non-relatives. Independent variables (Appendix A) were entered into the model in sequential groups of host, family environment, and then child care characteristics and were retained in the model using backward elimination if their p values were $<.05$ (two-tailed).

The suitability of linear regression for modeling non-normally distributed dependent variables (illness incidence) was validated by examination of the error terms. No serious deviation from normality was observed in these terms, and plots of residual terms and predicted values displayed no violation of the homoscedasticity assumption.[15,16]

Results

Cohort Members, Child Care Settings

Of 1,388 children under 36 months of age when the study began, 57 (4.1 per cent) were excluded due to chronic illness. Of the remaining 1,331, data were obtained by mail for 758 (57 per cent) and by telephone for an additional 85 (6 per cent). As there was no difference of consequence in the distribution of child care settings between these two groups of respondents (Appendix B), we combined them to yield a total of 843 children (63 per cent of the target population) upon whom our analysis is based.

Of the 843 children studied, the age distribution at entry was 33 per cent 0–11 months, 34 per cent 12–23 months, and 33 per cent 24–35 months; 48 per cent were male; and 57.5 per cent were White, 39.8 per cent Black, 0.1 per cent Hispanic, 1.2 per cent Asian, and 1.2 per cent of other or unspecified race. (In comparison, US Census data indicate that in 1980, the population of Shelby County, Tennessee, which includes Memphis, was 57.3 per cent White, 41.8 per cent Black, and 0.9 per cent of other or unspecified race.) Eighty-two per cent of the children studied lived in households with two adults, 11 per cent lived with only a single adult, and 7 per cent lived with three or more adults. Of the 685 households containing two adults, both adults were employed in 68 per cent.

The customary source of child care varied widely (Table 1), as did host and family environment characteristics of children in different forms of care (Appendix C). Compared with children enrolled in day care centers or mother's day out programs, those enrolled in day care homes were more likely to be permitted to attend if they were febrile (134/185 vs 43/200; odds ratio 9.6; 95 per cent CI = 5.9, 15.8), had symptoms of respiratory illness (162/185 vs 139/200; OR 3.1; 95 per cent CI = 1.8, 5.4), or had vomiting or diarrhea (102/185 vs 38/200; OR 5.2; 95 per cent CI = 3.2, 8.6).

Incidence of Illness and Associated Risk Factors

Bivariate analysis disclosed differences in infection rates according to child care setting (Table 2) which persisted after

TABLE 1—Child Care Arrangements of Study Participants

Child Care Setting	Number of Children Studied (%)	Mean Number of Hours Week in Child Care	Mean Number of Other Children in Room	Mean Cost per Week (Dollars)
In own home by parent	173 (20)	1.9	3.1	4.12
In own home by sitter	7 (1)	45.0	0.1	48.81
Relative's home	141 (17)	34.2	1.7	23.85
Day care home	194 (23)	42.9	3.4	35.28
Mother's day out	51 (6)	17.5	7.2	20.54
Day care center	159 (19)	42.3	13.0	43.96
Multiple settings	118 (14)	33.2	6.8*	27.75
All study participants	843 (100)	29.9	6.1	32.17

TABLE 2—Mean Number of Infections per Child Diagnosed during Study Period and Reported by Parents during One Month of Recall

Type of Child Care	Diagnosed by Physician in Prepaid Plan						Reported by Parent	
	Upper Respiratory Infection without Otitis Media	Lower Respiratory Infection	Otitis Media	Gastroenteritis	Skin Infection	Total Infections	Illnesses without a Physician Visit, in One Month of Recall by Parent	Illness Resulting in a Physician Visit outside Prepaid Plan, in One Month of Recall by Parent
Own home	0.79	0.16	0.84	0.15	0.09	2.03	0.66	0.28
Relative's home	0.68	0.18	0.77	0.09	0.10	1.82	1.00	0.29
Day care home	0.88	0.18	1.21	0.17	0.09	2.54	0.82	0.22
Mother's day out	1.18	0.20	1.49	0.10	0.18	3.14	0.90	0.16
Day care center	1.03	0.29	1.40	0.10	0.12	2.96	1.03	0.22
Multiple settings	1.13	0.17	1.37	0.21	0.12	3.00	0.99	0.38
All participants	0.91	0.19	1.13	0.14	0.11	2.49	0.90	0.28
P (F test, df = 5)	0.014	0.338	0.0006	0.147	0.723	0.0001	0.18	<0.0001

adjustment for potentially confounding host and family environment variables (Table 3). Hospitalization rates were unaffected by age and were higher for children in day care centers than for children in all other settings combined (relative risk = 4.50, 95% CI = 1.55, 13.00), primarily due to an increased incidence of tympanostomy tube placement (relative risk = 3.79, 95% CI = 1.04, 13.36) (Table 4).

Multivariate analysis identified the following predictors* of the total number of infections diagnosed: age in months (B = −.07, SEB = .01), breastfeeding prior to the study (B = .86, SEB = .24), median household income of census tract in $1000's (B = .094, SEB = .021), number of other children in the room in child care (B = .05, SEB = .02), lack of a separate room for diaper changing/potty use (B = .76, SEB = .28), and food furnished by the child care provider (B = .81, SEB = .37). Results were similar for individual categories of infection except that no child-care variables were predictive of upper respiratory infection without otitis media and presence of infants less than 12 months of age, acceptance of children with vomiting or diarrhea, and bringing food from home were predictive of gastroenteritis.*

No risk factors for skin infection were identified. Host, family environment and child care variables, respectively, accounted for 10, 4 and 5 per cent of the variance (R^2) in total number of infections diagnosed. Entry of an additional set variable, "child care setting", into the multivariate model resulted in a net increase in explained variance of only 0.7 per cent, suggesting that most of the variance attributable to child care was accounted for by variables specified in our model.

*Complete data from multivariate analyses are available from authors on request. B = unstandardized regression coefficient, SEB = standard error of B.

Financial Costs

The mean monthly cost of medical care, including prescription drug costs, for children in the highest risk settings (day

TABLE 3—Mean Number of Excess Episodes of Diagnosed Infection per Child for Children Receiving Care Outside of Their Own Homes, as Compared with Those Cared For in Their Own Homes, Adjusted for Host and Family Environment Variables (Appendix C)

Child Care Setting	Upper Respiratory Infection without Otitis Media	Lower Respiratory Infection	Otitis Media	Gastroenteritis	Skin Infection	Total Infections
	MEAN ACTUAL NUMBER OF EPISODES					
Own home (referent)	79	16	84	15	09	2.03
	MEAN EXCESS EPISODES (95% CI)					
Relative's home	−.10 (−.42, .21)	.03 (−.10, .15)	.08 (−.35, .50)	−.07 (−.19, .04)	−.02 (−.12, .07)	−.09 (−.73, .54)
Day care home	−.11 (−.41, .20)	−.05 (−.17, .07)	.31 (−.10, .72)	−.01 (−.12, .10)	−.06 (−.15, .03)	.09 (−.51, .71)
Day care center	−.10 (−.23, .43)	.11 (−.02, .24)	.66* (.22, 1.10)	−.09 (−.22, .03)	−.02 (−.11, .08)	.79* (.13, 1.45)
Mother's day out	−.15 (−.28, .57)	.04 (−.13, .20)	.44 (−.13, 1.01)	−.07 (−.22, .09)	−.06 (−.06, .19)	.60 (−.24, 1.46)
Multiple settings	−.15 (−.19, .48)	.02 (−.11, .15)	.51* (.06, .96)	−.04 (−.09, .16)	−.04 (−.14, .06)	.66 (−.01, 1.34)

*Significant excess compared with referent group (95% CI does not include 0.0).

TABLE 4—Incidence of Hospitalization during Study Period

Child Care Setting	Per Cent of Children Hospitalized	Reason of Hospitalization (number of patients)
Own home	1.7	Tympanostomy tubes (2) Croup (1)
Relative's home	0.0	—
Day care home	2.1	Tympanostomy tubes (3) Croup (1)
Mother's day out	2.0	Tympanostomy tubes (1)
Day care center	5.7	Hemophilus influenzae meningitis (1) Bronchiolitis (1) Gastroenteritis (1) Tympanostomy tubes (6)
Multiple settings	0.8	Tympanostomy tubes (1)
All participants	2.1	

care centers, mother's day out programs, and multiple child care settings) was $32.94, compared with $19.78 for other children (Table 5). Parents reported being absent from work a mean of 0.40 days per month between January and March 1986, of which a mean of 0.16 days (40%) was due to illness in a child in this study. The mean number of days absent from work per month because of illness in a child in this study was 0.52 for children in day care centers compared with 0.37 for children in other forms of full time day care outside the home (i.e., day care homes or relatives' homes).

Discussion

Our finding of an increased incidence of illness in children attending day care centers, but not day care homes, is consistent both with previous reports[5–9] and with the results of our multivariate analysis, which suggest that the major risk factor for infection in a child care setting is the number of other children in the room.

Our study has certain potential limitations. First, illness ascertainment bias may have been introduced if children in certain forms of child care were more likely to seek medical attention when ill or if, despite the incentives of the prepaid plan, they were more likely to make an unauthorized visit to a physician outside of our group. Our questionnaire data for one month within the study period suggested that neither of these possibilities occurred. Since symptoms in young children may be non-specific, a strength of our study was that illnesses were diagnosed by a physician.

Second, information on child care settings came from parents rather than from day care providers. We believed that the answers to our questions would be readily known to the parents and that they would respond truthfully to their child's pediatrician. This impression was corroborated by the prompt, unhesitating responses obtained in the telephone survey and by the fact that the average completion rate for questionnaires returned by mail was over 95 per cent.

Third, some risk factors identified in our multivariate model may actually have been proxies for other variables that were not studied. For example, the curious observation that a history of breastfeeding in the past was associated with a higher incidence of both total infections and upper respiratory infection without otitis media might be explained if breastfeeding, or its discontinuation, was a proxy for lack of immunologic experience or for a constellation of sociocultural practices that we could not better identify, but which led to an increased infection rate in the preschool years.

Finally, the financial data reported in this study pertain to families in Memphis, Tennessee enrolled in a prepaid health plan. These figures may not be applicable to other areas of the United States where medical charges may be different or to children covered by other types of health insurance in which there may be financial barriers to seeking medical care or different practicing styles of participating physicians. Our financial data, like our illness incidence data, are most useful in offering estimates of the magnitude of differences among children in different child care settings within a single study population. Our data indicate that there are certain potentially modifiable characteristics of day care which are risk factors for illness and that excess illnesses result in excess financial costs. The health of children in day care should be of concern not only to parents, physicians, and public health officials, but also to employers and health insurors who might find it worthwhile to subsidize the possibly increased cost of child care in settings designed to minimize illness rates.

ACKNOWLEDGMENTS

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TABLE 5—Mean Monthly Cost of Illness per Child in Study Participants (Dollars)

Type of Care	Physician's Office Visits	Emergency Room Visits	Prescription Drugs	Hospitalization	Total Medical Charges	Lost Parental Income* (Days per Month Absent from Work†)	Total Cost
In own home	12.74	0.08	4.11	2.51	19.45	9.17 (0.19)	29.50
Relative's home	11.19	0.77	3.52	0.00	15.48	16.05 (0.31)	31.22
Day care home	16.40	0.41	5.07	1.35	23.22	25.39 (0.42)	48.64
Mother's day out	21.23	0.07	6.34	1.90	29.54	12.21 (0.23)	39.75
Day care center	19.72	0.42	5.90	10.32	35.63	33.94 (0.52)	61.64
Multiple settings	19.16	0.31	5.95	4.31	30.78	26.61 (0.42)	58.30
All participants (N = 843)	16.05	0.37	4.96	3.52	24.90	23.01 (0.38)	46.68

*Reported for 651 children for whom the income of the parent staying home from work was known.
†To care for an ill child in this study.

APPENDIX A
Variables Entered into the Multivariate Models for Incidence of Infection

Host Variables

Age (months)
Sex
Breastfeeding during study
Breast-fed prior to study
Length of time (months) in child care prior to study

Environmental Variables

Median household income of census tract of residence
Sibling in day care
Number of household members in the age groups
< 12 months, 12 months–4 years
Household crowding index (people per room)
Number of smokers in household

Child Care Variables

Hours per week in child care
Number of other children in room
Age group(s) of other children in room
(< 12 months, 12 months–4 years)
Children caregiver ratio in child's room
"Drop-in" children permitted
Child accepted with: 1) respiratory symptoms
2) vomiting or diarrhea
Food furnished by child care provider
Food brought from home
Separate room for diaper changing, "potty," and toilet use
Caregivers known to smoke in room with children
Cost per hour of child care

Child Care Settings

Day care home
Mother's day out program*
Day care center
Multiple settings

*A form of group care common in the Memphis area, usually sponsored by a church one or two days per week.

APPENDIX B
Comparison of Respondents and Non-Respondents

Data from Questionnaires	Questionnaire Returned by Mail N = 758	Questionnaire Administered by Telephone N = 85	Non-respondents N = 488
Child Care Setting			
% In own home	21	26	N.A.
% Relative's home	16	20	N.A.
% Mother's day out	6	4	N.A.
% Day care home	23	24	N.A.
% Day care center	19	18	N.A.
% Multiple settings	15	7	N.A.
Mean age (mos)	17.6	17.4	18.8
% White	58	46	N.A.
Median household income of census tract of residence	$20,291	$18,412	$17,797
Mean no. of prepaid plan visits	4.40	3.14	3.09
Mean no. of prepaid plan visits for infection	2.57	1.87	1.77
Mean no. of prepaid plan visits for well child care	0.82	0.61	0.57
Mean no. times child sick but not seen by physician in one month of recall by parent	0.80	0.91	N.A.
Mean no. visits to emergency room or outside prepaid plan in one month of recall by parent	0.29	0.13	N.A.
Mean no. days absent from work by parent January–March 1986	2.5	3.7	N.A.
Mean no. days absent to care for ill index child, January–March 1986	1.1	1.5	N.A.
Income of parent staying home	$10,750	$11,650	N.A.

N.A. Not available.

APPENDIX C
Host and Family Environment Characteristics which Varied Significantly among Children in Different Child Care Settings

Setting	Mean Number of Respondents*	Mean Age (months)	Per Cent White	Median Income of Census Tract of Residence	Per Cent of Infants under 12 Mos of Age Breast-fed Throughout Study	Per Cent of All Participants Who Had Been Breast-fed Prior to Study
Own home	171	15.4	53	$19,494	9	51
Relative's home	136	16.5	31	$18,068	1	40
Day care home	189	14.4	58	$20,110	2	52
Mother's day out	50	19.7	85	$22,769	2	70
Day care center	149	22.9	68	$20,513	1	55
Multiple settings	112	18.6	68	$20,986	4	52
All participants	813	17.6	57	$20,104	3	52

Setting	Mean Time in Child Care Prior to Study (months)	Mean Number of Household Members	Per Cent of Children with ≥1 Sibling in Child Care	Mean Household Crowding Index (persons/room)	Mean Number of Smokers in Household
Own home	1.0	4.4	11	0.75	0.6
Relative's home	9.2	3.9	15	0.70	0.8
Day care home	9.8	3.7	51	0.59	0.7
Mother's day out	5.9	3.8	20	0.81	0.3
Day care center	14.3	3.6	20	0.64	0.5
Multiple settings	11.1	4.3	34	0.66	0.6
All participants	8.7	3.9	33	0.68	0.6

*N varies slightly among cells within each row.

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NCHS Quarterly Report Cites New Publications, Findings

The National Center for Health Statistics recently announced, in its Quarterly Report, several new publications to be available from the Center in February 1989. These include the following:

- *Characteristics of Nursing and Related Care Homes: The 1986 Inventory of Long-Term Care Places*. (Describes facility characteristics such as ownership, certification status, numbers of beds and residents. Examines occupancy rates, per aged population, and numbers of Black and Hispanic residents.)
- *The National Nursing Home Survey: 1985 Summary for the United States*. (Provides descriptive tables on characteristics of nursing homes, registered nurses at the facilities, current residents, data on cost of providing care, health and functional status of residents, payment for care.)
- *Children of Divorce*. (Analyzes trends in the number and proportion of children whose parents are divorced, as well as characteristics of divorcing couples with and without children, and geographic variations.)
- Also, Coming Soon: *NCHS 1988 Year in Review* (A new publication featuring two series of articles published during 1988 on NCHS data and programs. And an overview of surveys conducted and information available from NCHS for 1988.)

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Recurrent Outbreaks of Giardiasis in a Child Day Care Center, Wisconsin

Richard W. Steketee, MD, MPH, Steven Reid, MD, Tina Cheng, MD, James S. Stoebig, MS, Richard G. Harrington, MS, and Jeffrey P. Davis, MD

Abstract: In the 19-month period September 1983–March 1985, three outbreaks of giardiasis occurred in one large child day care center. Control measures instituted during each outbreak included case finding; pharmacologic treatment and follow-up testing of stool specimens for cases of giardia infection in day care children and staff, and their household contacts; facilitating and stressing personal and environmental hygiene, including altering diapering practices and teaching appropriate hand washing techniques. In the first, second, and third outbreaks, overall attack rates (stool analysis positive for *Giardia lamblia*) were determined for those persons with ≥2 stool specimens submitted; attack rates in children were 47, 17, and 37 per cent, respectively; for tested staff, the rates were 35, 13, and 9 per cent; and for tested household contacts were 18, 9, and 5 per cent. Attack rates were highest for ambulatory diapered children, children attending the day care center ≥40 hours per week, and children who were infected with *G. lamblia* in the most recent previous outbreak. Despite extensive efforts to identify cases, a cure rate of >90 per cent in treated cases, and improvements in personal and environmental hygiene practices, *G. lamblia* infections recurred in outbreak proportions. (*Am J Public Health* 1989; 79:485–490.)

Introduction

Gastrointestinal infections associated with child day care have been of increasing interest in the United States in recent years, and diarrheal illness caused by *Giardia lamblia* has been noted frequently.[1] Within the child day care setting, efforts to understand the occurrence of giardiasis have included assessing the incidence and the prevalence of *G. lamblia* infections,[2–6] the extent of symptomatic illness,[2,3,5] the modes of transmission,[2,4,7,8] the specific risk factors for acquisition of *G. lamblia* infection,[2,5,6] and the appropriate control measures.[2,5,9–11]

In September 1983, an outbreak of symptomatic giardiasis was detected in a proprietary child day care center in Wisconsin. Control measures included: case finding in symptomatic and asymptomatic children, staff, and household contacts of infected children; pharmacologic treatment and follow-up test-of-cure of all cases of *Giardia* infection; and health education for staff, including regular teaching of personal and environmental hygiene. A second outbreak of giardiasis was detected in this day care center within eight months of the first outbreak and a third outbreak occurred eight months after the second outbreak. We present here a description and evaluation of the giardiasis outbreak control measures in this day care setting.

Background

During the interval from September 1983 to March 1985, this day care center in suburban Dane County, Wisconsin provided care for children ages 1 month to 6 years and had an average daily census of 115 children of whom 45 per cent were toilet trained. The facility is licensed for the care of 175 children and the enrollment ranged from 100 to 167 children during the 18-month interval. The fluctuation in enrollment was due largely to changes in the number of preschool and kindergarten children, while the number of infants and toddlers remained relatively stable. Children came from predominantly middle-income households in neighborhoods surrounding the center. The average child-staff ratio was 7.6 to 1 (range 6.6 - 10.4 to 1), the center provided an average of 59 square feet of classroom space per child (range 40.6–60.6 square feet per child). Play groups were organized by age groups as shown in Table 1 with 10 to 15 children within a play group; each play group was located in a separate classroom. Mixing of play groups occurred occasionally during outdoor playtime, but was otherwise uncommon. The day care center staff worked only with their assigned play group and had minimal contact with other play groups.

The center's staff maintained daily records of bowel movements for children ≤36 months old and the stool consistency was noted to be formed, loose, or watery; the parent was informed if the child had a perceived change in bowel habits or watery diarrhea. The day care center's management policy for children with diarrhea included notifying parents and strongly encouraging exclusion of symptomatic children. Although compliance with this policy was not routinely documented, staff felt that parents nearly always accepted the policy. Due to limitations of space and staff supervision, cohorting of ill children was not practiced. At the time of the initial outbreak investigation, food handlers also provided child care. Diaper changing occurred in the classrooms and, except in the infant rooms, running water was available only in bathrooms. A bucket of water was used for washing the diapering table and for hand washing after diapering. The water bucket was emptied and refilled between diaper changes.

Methods

Case Finding

On September 2, 1983, the director of the day care center reported that four children in one play group had developed a diarrheal illness, and that 33 per cent of 87 infants and toddlers had experienced diarrhea in the previous two weeks. After obtaining parental consent, stool specimens for bacterial and parasitic pathogen testing were collected from these four children by the local public health agency; *G. lamblia* was the only enteric pathogen identified in their stool specimens. With support from the local public health agency, efforts were made to collect two or more stool specimens from all symptomatic day care center attendees and six

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TABLE 1—Giardiasis Attack Rates (AR), by Age Play Group, Day Care Center (DCC) Outbreaks A, B, and C, Dane County, Wisconsin, September 1983–January 1985

Play Group	Age (months)	Outbreak A Total n	Outbreak A Cases Tested† n	Outbreak A AR	Outbreak B Total n	Outbreak B Cases Tested n	Outbreak B AR	Outbreak C Total n	Outbreak C Cases Tested n	Outbreak C AR
Infants	0–17	20	9/18	50	31	2/26	8	24	0/19	0
Toddlers 1	18–23	14	7/13	54	15	4/15	27	17	6/17	35
Toddlers 2	24–29	17	8/16	50	14	3/10	30	20	8/14	57
Toddlers 3	30–35	7	3/4	75	21	1/14	7	20	1/5	20
Preschool*	36–72	57	14/36	39	25	2/4	50	86	3/9	33
Total Children		115	41/87	47	106	12/69	17	167	18/64	28
Staff		17	6/16	38	16	1/8	13	16	1/11	9
Household Contacts‡ of DCC cases		61	8/47	17	19	1/11	9	28	1/21	5
Total	—	193	55/150	37	141	14/88	16	211	20/96	21

*Testing in preschool groups during outbreak B and C included only symptomatic children and children who had DCC attending siblings with cases giardiasis.
‡Household contacts included current permanent residents in households with a DCC case of giardiasis.
†Tested with ≥2 stool specimens.

additional persons with *G. lamblia* infections were identified. Stool specimens were then collected from asymptomatic day care attendees in the same play groups as known cases. Investigation and public health measures were instituted and this outbreak (outbreak A, see definition below) ended in late October 1983. Two additional outbreaks of giardiasis were identified in this center in the following 17 months: outbreak B occurred in June–July 1984, and outbreak C occurred in January–March 1985. Each outbreak was identified and evaluated in the same manner by local and state public health officials.

Definitions

- A *case* of giardiasis was defined as infection (with or without symptoms of diarrhea) in any attendee, staff, or household contact with laboratory confirmed *G. lamblia* in ≥1 stool specimen.
- A *non-case* was defined as person with ≥2 stool specimens collected more than 24 hours apart that were negative for *G. lamblia*.
- *Diarrhea* was defined as an increased frequency of stools with a change in consistency of stools to watery or loose as reported by parents or day care center staff. Center staff used this definition in routine surveillance for diarrhea in the day care center.
- An *outbreak* of giardiasis was defined as an increased prevalence of diarrheal illness among day care attendees with laboratory-confirmed *G. lamblia* infection in >20 per cent of screened attendees (with or without symptoms) in any age or play group. Although staff maintained records of day care attendee bowel movements during the day, these records were not retained more than two weeks. Thus, the "increased prevalence of diarrhea" in this definition relied on staff perception of an increase.
- An *outbreak* was classified as ended when the prevalence of diarrhea in day care attendees was reduced to "pre-outbreak levels" and no new cases were identified within a two-week period. Again, "pre-outbreak levels" relied on staff perception, but could be verified as lower than the point prevalence of diarrhea during the outbreak by reviewing bowel movement records which were maintained throughout the outbreak.
- The *timing of outbreak* onset and termination was established based on the combined requirement of increased prevalence of diarrheal signs and symptoms coupled with laboratory results, thus dates for each of the outbreaks are given by month rather than day or week of onset and termination.

Control Measures

Control measures used during the outbreaks included systematic case identification among day care attendees, staff, and household contacts of cases, using the algorithm in Figure 1. Persons with confirmed giardiasis were treated for 7–10 days by their physician with quinicrine hydrochloride, furazolidone, or metronidazole.* Parents of a case-child were asked to keep the child at home until the diarrhea resolved. Three weeks after completion of treatment, two test-of-cure stool specimens obtained >24 hours apart were examined for *G. lamblia* cysts. Retreatment was instituted as necessary until two stool specimens collected >24 hours apart were negative. Staff-cases were treated in similar fashion with medication, excluded while symptomatic, and given test-of-cure follow-up. Household-cases were treated in similar fashion with medication and the test-of-cure follow-up.

During outbreak A, changes were made in the center's routine to interrupt *Giardia* transmission. A public health nurse met with the staff of the day care facility to provide education regarding the signs, symptoms, and methods of diagnosis of *Giardia* infection, and the methods of prevention or interruption of transmission including personal and environmental hygiene. A public health nurse made follow-up visits to the center several times each week during the outbreak and at least monthly during the non-outbreak periods. Within the center, the diapering areas were moved closer to the sinks to facilitate hand washing and cleansing of diapering tables. Diapering areas were supplied with paper to be placed under each child during diapering and discarded after each diaper change. Food handlers were excluded from child care responsibilities. Appropriate hand washing procedures were taught and stressed by a public health nurse during follow-up visits and written guidelines for handwash-

*Names of drugs are for identification purposes only and do not imply endorsement by the US Department of Health and Human Services or US Public Health Service.

FIGURE 1—Guidelines Used for Investigation and Pharmacologic Intervention in Day Care Center (DCC) Outbreak of Giardiasis, Wisconsin

ing practice were posted in the day care center. No change was made in the practice of monthly washing of all day care center toys.

The control measures used in outbreak A were also employed in outbreaks B and C, except that testing in children ≥3 years old was limited to symptomatic children or siblings of known *Giardia*-positive children in the center.

Investigation Methods

Information collected from the center's records for each attendee included name, age, gender, toilet training status, number of hours per week at the center, and date of initial enrollment. The day care staff provided information on the child's gastrointestinal signs and symptoms during the previous four weeks. During outbreaks B and C, we reviewed our records from the previous outbreak to determine the case status of the child during the most recent previous outbreak.

Laboratory Methods

In each outbreak, stool specimens were obtained from the initial symptomatic day care attendees and were tested for bacterial enteric pathogens (*Salmonella*, *Shigella*, *Campylobacter* and *Yersinia* species) and for ova and parasites; testing for enteric viruses was not done since routine stool testing in this setting was designed to identify pathogens which require specific antimicrobial therapy. In each outbreak, *G. lamblia* was the only pathogen detected and subsequent screening was for *Giardia* only. Two or three stool specimens were collected >24 hours apart in the day care center or at home and placed in a vial containing formalin. The specimens were centrifuged and the concentrate was examined for 3–5 minutes by light microscopy by a parasitologist. Specimens were examined both unstained and stained with trichrome stain.[12] If no *G. lamblia* cysts were seen, the stool concentrate was reconstituted in formalin-ether,[13] centrifuged a second time, and re-examined in the same fashion. Stools were considered positive if any *G. lamblia* cysts were seen. Test-of-cure stool specimens were examined in the same fashion.

Data Analysis

Comparisons of attack rates in different groups are presented as incident rate ratios (IRR) with 95% confidence intervals (CI).[14]

Results

Table 1 summarizes the *G. lamblia* attack rates for the three outbreaks. In outbreaks A, B, and C, *G. lamblia* was identified in stools of 47 per cent, 17 per cent, and 37 per cent of tested children; 38, 13, and 9 per cent of tested staff, and in 17, 9, and 5 per cent of tested household contacts, respectively. Among day care center children, *Giardia* infections were most common in Toddler Groups 1 and 2 (ages 18–23 months [38 per cent positive] and 24–29 months [48 per cent positive]), and were least common (17 per cent) in the infant group.

The *Giardia* attack rate (43 per cent) for 75 children attending day care ≥40 hours per week was significantly greater than the rate of 10.5 per cent for 137 children attending <40 hours per week, (IRR = 4.17, 95% CI 2.17, 8.47); this association was seen in all age groups.

Completed questionnaires were obtained for a total of 217 children with known symptoms and case status. Diarrhea within four weeks of recognition of the outbreak was reported in 35 (51 per cent) of 68 *G. lamblia* infected children and 42 (28 per cent) of 149 non-case children (IRR = 1.83, 95% CI 1.13, 2.93). Case children were more likely than non-case children to report symptoms of abdominal cramps, 13 percent versus 4 per cent, respectively (IRR = 3.29, 95% CI 1.05, 11.22). No significant differences between case children and non-case children were observed for other reported symptoms (respectively, vomiting in 0% and 1%; fever in 1% and 0%; loss of appetite in 12% and 5%; and "gas" in 9% and 3%). Among the *Giardia* infected children, 49 per cent were asymptomatic.

Information on the use of drug treatment was available for 89 case-persons (children and staff) and 81 (91 per cent) were treated with medication. Two negative test-of-cure stool specimens were obtained for 52 per cent of treated individuals and an additional 33 per cent had one test-of-cure negative exam. Table 2 shows treatment results for 67 persons for whom the specific medication and dose was known. The three drugs (metronidazole, furazolidone, and quinacrine hydrochloride) produced an overall >90 per cent cure rate in the 64 persons who had ≥1 test-of-cure stool examination.

At the time of detection of outbreak B, 37 (54 per cent) of the 69 tested day care children were newly enrolled after outbreak A, and at the time of outbreak C, 29 (45 per cent) of the 64 tested day care children were newly enrolled after outbreak B. During the two months of the outbreak A investigation, only 77 (57 per cent) of the center children attended throughout the outbreak, while 28 (21 per cent) enrolled during the outbreak and 29 (22 per cent) withdrew from the center during the outbreak. During outbreak A, attack rates were similar for those who stayed throughout and those who withdrew (52 and 50 per cent, respectively) and were lower in the newly enrolled children (25 per cent).

Attack rates by case status during the previous outbreak are shown for outbreaks B and C in Table 3. The day care center children in whom cases occurred during the previous outbreak were more likely to be case individuals than

TABLE 2—Pharmacologic Treatment and Follow-up of 72 Cases of Giardiasis, Day Care Center outbreaks A, B, and C, Wisconsin, August 1983–March 1985

Outbreaks	Number of Cases	Drug*	Number Treated	Follow-up Stool Test	Cure Rate† n	Cure Rate† (%)
A	41	M	6	6	6	(100)
		F	20	20	19	(95)
		Q	15	15	14	(93)
B	11	M	1	1	1	(100)
		F	9	8	7	(88)
		Q	1	—	—	
C	20	M	4	3	3	(100)
		F	9	9	7	(78)
		Q	2	2	2	(100)
Total	72	M	11	10	10	(100)
		F	38	37	33	(89)
		Q	18	17	16	(94)

*M = metronidazole, F = furazolidone, Q = quinicrine hydrochloride.
†Two or more follow-up stool specimens were available on 46 (69%) persons and an additional 18 (26%) had only one follow-up test. All failures were identified on first follow-up specimen.

children who were previously not cases individuals or children who were new arrivals into the center. Eight of the center children had *Giardia* infection during two consecutive outbreaks. Six of these children had ≥2 negative test-of-cure stool specimens, one child had one negative test-of-cure specimen, and one child had no post-treatment follow-up between the first and second infection. At the time of their second documented infection, four children had diarrhea and four were asymptomatic; this is no different from the rate of symptomatic *Giardia* infection in previously negative children or new enrollees. The one staff person with recurrent infection had three negative stool specimens between her first and second infection; she was asymptomatic during both infections.

Compliance with the personal and environmental hygiene control measures was not measured systematically; however, public health nurses observed proper handwashing techniques during their follow-up visits and diapering practices were altered during outbreak A and maintained thereafter as described above.

Discussion

Following efforts to control a day care center outbreak of giardiasis including extensive case-finding, high rates and success of pharmacologic treatment of cases, and changes in personal and environmental hygiene practices, *Giardia* outbreaks were observed to end but subsequently recurred on two occasions within eight months of the preceding outbreak. Possible impediments to effective control efforts included the new enrollment of infected children into the center, the continued shedding of *Giardia* by children or staff awaiting treatment, and environmental or design limitations of the facility which required staff to use buckets of water for handwashing and cleaning the diaper changing table. Although we cannot implicate a specific problem with the control measures used and we did observe a decrease in the prevalence of diarrhea and *G. lamblia* infection after instituting the measures, our findings suggest that treating all identified *G. lamblia* cyst passers in conjunction with alterations in handwashing and diaper changing practices did not permanently eliminate giardiasis transmission in this day care facility.

During these investigations, the designation of an outbreak of giardiasis in the day care center required a staff-perceived increase in diarrheal symptoms and laboratory confirmed *G. lamblia* in >20 per cent of day care children in one or more play groups. Only symptomatic children and their play group contacts were tested; the overall prevalence of infection in the center was lower than 20 per cent. Although studies in the southern US[4] have shown point prevalence rates of >20 per cent in day care centers without staff-perceived increased diarrhea prevalence, this has not been observed in Wisconsin (unpublished findings) or to our knowledge elsewhere in the northern US. Consequently, our definition was used as a guideline for public health intervention in Wisconsin and may not be applicable elsewhere in the country.

The determination of an outbreak may also have been limited by the lack of testing for a variety of viral enteric pathogens. Diarrhea within the previous four weeks was reported in 28 per cent of children without *G. lamblia* infection. It is possible that the some of the increase in diarrheal stools in day care attendees was due to a viral pathogen(s) or other etiologic agents and that *G. lamblia* was found as an incidental organism in some of these children. Regardless of the potential presence of additional pathogens, *G. lamblia* was detected in a high proportion of children with

TABLE 3—Giardiasis Attack Rates (AR), for Day Care Center Attendees in Outbreaks B and C, by *Giardia* Case Status During Prior* Outbreak

Outbreaks	New children enrolled since prior outbreak: Number	Cases in new children: No.	Cases in new children: (AR)	Children with case in prior outbreak: Number	case this outbreak: No.	case this outbreak: (AR)	Children with no case in prior outbreak: Number	case this outbreak: No.	case this outbreak: (AR)
B	37	5	(14%)	14	5	(36%)	18	2	(11%)
C	29	4	(14%)	5	3	(60%)	30	11	(37%)
TOTAL	66	9	(14%)	19	8	(42%)	48	13	(31%)

The incident rate ratio (IRR) for comparison between groups is as follows:

Comparison	IRR	95% CI
Child previously present and positive vs new child	3.0	1.04, 9.02
Child previously present vs new child	2.30	1.01, 5.70
Child previously positive vs child previously negative	1.35	0.56, 4.05

*Prior outbreak refers to most recent outbreak preceding this outbreak. One child had a case of *Giardia* infection in all three outbreaks.

diarrheal illness and was the only detected pathogen amenable to antimicrobial therapy. Thus, the short-lived effect of antimicrobial therapy as evidenced by the resumption of high *G. lamblia* prevalence demonstrates the limited benefit of this treatment as a control measure.

Giardiasis control efforts during these outbreaks depended in part on prompt diagnosis and treatment of cases. *Giardia* case-finding is hampered by false-negative results of microscopic examination of fecal material,[15] the diagnostic method used in this investigation. Alternative methods including duodenal aspirates or biopsy examination and in vitro culture also show false-negative results and the complexity of methods precluded their use in this investigation.[16] Although heightened surveillance and screening during the outbreaks probably identified a majority of cases, only selected groups of persons had stool examinations and some cases may have gone undetected; persons with undetected cases may have continued to transmit *G. lamblia*. Also, transmission possibly occurred prior to institution of drug therapy for cases. Although drug treatment was instituted in >90 per cent of identified cases, completion of two test-of-cure follow-up stool examinations was achieved for only 52 per cent of cases. Those without post-treatment stool examinations would be expected to have a cure rate similar to the >90 per cent cure rate for tested persons, but some of these persons may have been still capable of transmitting infection. Several cases required retreatment due to continued shedding of *G. lamblia* cysts; transmission may have occurred during this interval of cyst shedding.

By the onset of outbreak B, 58 per cent of the attendees <36 months old were new to the day care facility and by the onset of outbreak C, 33 per cent of attendees <36 months old were new to the center. No systematic assessment employing the collection of multiple stool specimens has been made to determine the prevalence of *G. lamblia* infection in the Dane County child population, but using an estimate of 5 per cent prevalence derived from a review of *G. lamblia* surveys,[17] at least one or two of the entering children would be expected to have a *Giardia* infection. During the interval of the three outbreaks, four staff terminated and four new staff were hired (a 24 per cent turnover of staff). This is consistent with the approximately 30 per cent annual turnover rate among the day care employees observed in Dane County, Wisconsin (personal communication, Lynn Mussallem, RN, Madison City Health Department). Although staff turnover may not be as important in the introduction of *Giardia* infection and the continued education provided throughout the interval of outbreaks should have minimized the effect of staff turnover on the control measures, it is possible that the efforts toward consistent hand washing and environmental control practices were hindered by high staff turnover rates.

Attack rates were higher for children who had previous *Giardia* infections. Additionally, attack rates in the children were high in each of the outbreaks but decreased among the staff from outbreak to outbreak. One possible explanation for this observation is that previously positive children may have continued to harbor *G. lamblia* cysts despite test-of-cure negative stools. Alternatively, children with previous *G. lamblia* infection may have behavior patterns which increase their exposure to the organism. Acquired immunity to *Giardia* infection has been suggested in adult human populations[18–20] and there is experimental evidence of acquired immunity in mice.[21,22] It is possible that immunity is not established in young children or that immune response is blunted by early drug therapy, but these factors could not be assessed in our investigation. Conceivably, a child's behavior which fosters *Giardia* transmission puts that child at increased risk for continual exposure and, despite even partial immunity, recurrent infection may be common in these individuals.

The most effective means for control of *Giardia* transmission in the day care setting is still debated. Some investigators have suggested treating all day care-associated cases of giardiasis regardless of symptoms because of the possibility of transmission of *G. lamblia* by asymptomatic cyst-passers[2,11] while others have encouraged treatment only in symptomatic children.[5,10] During these three outbreaks in Wisconsin, 93 per cent of *Giardia*-infected children and staff and all 10 households contact cases were pharmacologically treated regardless of symptoms. Day care center children with previous *Giardia* infection may be at increased risk for re-infection or recrudescence of old infection and should receive special follow-up. Management policies which minimize high turnover rates among employees and attendees may be beneficial in the control of *G. lamblia* transmission and possibly other infections. A carefully conceived plan for environmental preventive measures and drug therapy of symptomatic children, symptomatic household contacts, and possibly certain asymptomatic *G. lamblia* cyst passers should be established in child day care facilities. Our findings suggest that, given the ubiquity of the *G. lamblia* organism and because of the time commitment required for case identification/treatment/follow-up and the potential for drug treatment failures, hygienic control measures should receive increased emphasis in day care settings. However, clinicians, public health workers, and day care personnel can expect day care center outbreaks of giardiasis to recur.

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National Conference on HBP Control, May 6–9 at Disney World Village

The National Conference on High Blood Pressure Control, to be held May 6–9, 1989 at the Hilton Hotel, Walt Disney World Village in Florida, will address many of today's concerns for HBP control in special populations. High blood pressure continues to be the major cause of stroke and kidney failure and a principal contributor to heart attack deaths—costing the nation hundreds of thousands of lives and billions of dollars each year.

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A New Epidemiologic and Laboratory Classification System for Paralytic Poliomyelitis Cases

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Abstract: An epidemiologic classification of paralytic poliomyelitis cases (ECPPC) has been in use in the United States since 1976. In 1985, this classification system was reviewed because of recent changes in the epidemiology of paralytic poliomyelitis and improved laboratory capability to definitively characterize poliovirus strains. An alternative classification system was devised, the epidemiologic and laboratory classification of paralytic polio cases (ELCPPC), that incorporated virus isolation and strain characterization with epidemiologic information. Reported paralytic poliomyelitis cases for 1980–86 were classified by both the ECPPC and the ELCPPC classification systems. The new ELCPPC system classified 91 per cent of the reported cases as vaccine-associated, while the ECPPC system classified only 71 per cent of the reported cases as vaccine-associated. The proposed classification system provides more specific and useful information particularly concerning vaccine-associated paralytic poliomyelitis. (*Am J Public Health* 1989; 79:495–498.)

Introduction

Poliomyelitis control efforts have been very successful in the United States. Paralytic poliomyelitis cases declined from approximately 16,000 per year in the early 1950s, the immediate pre-vaccine era, to an annual average of less than 10 cases during 1980–86.[1,2]

A system for classification of paralytic poliomyelitis cases known as "epidemiologic classification of paralytic poliomyelitis cases" (ECPPC) has been in use since 1976.[3] Both the last epidemic of poliomyelitis and the last known case of endemic paralytic poliomyelitis due to wild poliovirus in the United States occurred in 1979.[4] Approximately one case of imported paralytic poliomyelitis is reported annually. Oral polio vaccine (OPV), in wide use since 1960, has been associated with a small number of cases of paralytic poliomyelitis annually among vaccine recipients and their susceptible contacts.[5,6] The overall risk of vaccine-associated paralytic poliomyelitis for the period 1973–84 was estimated as one case of paralytic poliomyelitis per 2.6 million doses of OPV distributed.[6] With the elimination of epidemic and endemic cases of paralytic poliomyelitis and only one imported case per year, vaccine-associated paralytic poliomyelitis has become more important. Poliovirus isolates were first reliably identified at the Centers for Disease Control (CDC) by their antigenic properties, using specific polyclonal antibodies,[7] an approach further refined by the use of cross-adsorbed antisera described by van Wezel.[8] While the specificity of antigenic characterization has improved with the development of panels of neutralizing monoclonal antibodies,[9–12] polioviruses are most definitively identified by molecular methods such as oligonucleotide fingerprinting or partial genomic sequencing,[13,14] which measure the extent of genetic sequence relationship among isolates. Molecular characterizations, used at CDC since 1980, have generally (> 90 per cent) confirmed identifications made on the basis of antigenic properties. Confirmation of suspected paralytic poliomyelitis cases depends on medical, epidemiological, and laboratory data. The isolation of polioviruses in stool specimens of patients with suspected poliomyelitis contributes to the decision-making process but does not constitute proof of a causal association of such viruses with paralytic poliomyelitis.

We describe a new classification system for reported paralytic poliomyelitis cases which integrates epidemiologic criteria, virus isolation, and strain characterization results, and then compare it with the existing classification system by applying both systems to reported paralytic poliomyelitis cases in the United States for the period 1980–86.

ECPPC Classification System

The ECPPC classification system distributes all cases into one of four categories: epidemic, endemic, imported, or immune deficient (Table 1). The system classifies cases based on epidemiologic information alone. When it was established, laboratory tests were not available for definitively differentiating between strains of vaccine-like and wild polioviruses.

ELCPPC Classification System

In 1985, expert consultants were asked by CDC to review the applicability of the ECPPC system because of

TABLE 1—Epidemiologic Classification of Paralytic Poliomyelitis Cases (ECPPC): United States, 1980 through 1986

	Categories	Number of Cases
I.	EPIDEMIC	
	A. Oral Polio Vaccine (OPV) not received from 4–30 days before onset of illness	0
	B. OPV received from 4–30 days before onset of illness	0
II.	ENDEMIC	
	A. No history of receiving OPV or of contact with an OPV recipient as defined in B and C below	6
	B. OPV received from 4–30 days before onset of illness	23
	C. Onset of illness 4–60 days after OPV was fed to a recipient in contact with the patient *and* contact occurred within 30 days before onset of illness	
	1. Household contact—vaccinee and patient regularly share the same home for sleeping	14
	2. Community contact or non-household contact	10
III.	IMPORTED—Disease develops in United States resident who has traveled outside the United States in areas with known endemic or epidemic poliomyelitis	5
IV.	IMMUNE DEFICIENT	8
	TOTAL	66

From the Centers for Disease Control, Atlanta. Drs. Sutter, Brink, Cochi, Orenstein, and Biellik are with the Division of Immunization, Center for Prevention Services. Dr. Kew is with the Division of Viral Diseases, Center for Infectious Diseases. Dr. Hinman is with the Center for Prevention Services, all are at CDC in Atlanta. Address reprint requests to Technical Information Service, Center for Prevention Services, CDC, Atlanta, GA 30333. This paper, submitted to the Journal August 8, 1988, was revised and accepted for publication November 2, 1988.

TABLE 2—Epidemiologic and Laboratory Classification of Paralytic Poliomyelitis Cases (ELCPPC): United States, 1980 through 1986

Categories	Number of Cases
I. SPORADIC	
A case of paralytic poliomyelitis not linked epidemiologically to another case of paralytic poliomyelitis	
A. Wild virus poliomyelitis. Virus characterized as wild virus	0
B. Vaccine-associated poliomyelitis	
1. Recipient—OPV was received 4 to 30 days before onset of illness	
a. Virus characterized as vaccine-related	16
b. Virus not isolated or not characterized	7
2. Contact—Illness onset was 4 to 75 days after OPV was fed to a recipient in contact with patient and contact occurred within 30 days before onset of illness	
a. Household—vaccinee and patient regularly share the same household for sleeping	
i. Virus characterized as vaccine-related	11
ii. Virus not isolated or not characterized	5
b. Non-household	
i. Virus characterized as vaccine-related	6
ii. Virus not isolated or not characterized	3
3. Community—No history of receiving OPV or of contact with an OPV recipient, as defined in 1 and 2, and virus isolated and characterized as vaccine-related	4
C. Poliomyelitis with no history of receiving OPV or of contact with an OPV recipient, as defined in B1 and B2, and virus not isolated or not characterized	1
II. EPIDEMIC	
A case of paralytic poliomyelitis linked epidemiologically to another case of paralytic poliomyelitis	
A. Not a recipient of OPV	
1. Virus characterized as wild virus	0
2. Virus characterized as vaccine-related	0
3. Virus not isolated or not characterized	0
B. OPV recipient—OPV received 4 to 30 days before onset of illness	
1. Virus characterized as wild virus	0
2. Virus characterized as vaccine-related	0
3. Virus not isolated or not characterized	0
III. IMMUNOLOGICALLY ABNORMAL	
Proven or presumed	
A. Wild virus poliomyelitis—Virus characterized as wild virus	0
B. Vaccine-associated poliomyelitis	
1. Recipient—OPV was received 4 to 30 days before onset of illness	
a. Virus characterized as vaccine-related	5
b. Virus not isolated or not characterized	0
2. Contact—Illness onset was 4 to 75 days after OPV was fed to a recipient in contact with patient and contact occurred within 30 days before onset of illness	
a. Household—vaccinee and patient regularly share the same household for sleeping	
i. Virus characterized as vaccine-related	1
ii. Virus not isolated or not characterized	0
b. Non-household	
i. Virus characterized as vaccine-related	1
ii. Virus not isolated or not characterized	0
3. Community—No history of receiving OPV or of contact with an OPV recipient, as defined in 1 and 2, and virus isolated and characterized as vaccine-related	1
C. Poliomyelitis with no history of receiving OPV or of contact with an OPV recipient, as defined in B1 and B2, and virus not isolated or not characterized	0
IV. IMPORTED	
Poliomyelitis in a person (US resident or other) who has entered the United States	
A. Virus characterized as wild virus	
1. Onset of illness within 30 days before entry	2
2. Onset of illness within 30 days after entry	1
B. Virus characterized as vaccine-related	
1. Onset of illness within 30 days before entry	0
2. Onset of illness within 30 days after entry	0
C. Indeterminate—Virus not isolated or characterized	
1. Onset of illness within 30 days before entry	1
2. Onset of illness within 30 days after entry	1
TOTAL	66

recent changes in the epidemiology of paralytic poliomyelitis in the United States and the improved laboratory capability since 1980 to definitively characterize poliovirus strains. The consultants were requested to:

- provide revisions of the system or an alternative if in their view the current classification system was no longer adequate;
- apply any revised or new system to cases of poliomyelitis with onset of illness from 1980–84 that met the CDC Paralytic Poliomyelitis case definition*;
- classify the same cases of poliomyelitis by the older ECPPC system.

For reported cases of paralytic poliomyelitis with onset in 1985 and 1986 and for cases with onset in 1980–84 but reported after the 1985 review, personnel at CDC assigned cases to classification categories in both systems. The consultants recommended that an alternative classification system be adopted that would incorporate virus isolation and strain characterization results and epidemiological information. The system, labeled the Epidemiologic and Laboratory Classification of Paralytic Poliomyelitis Cases (ELCPPC), is presented in Table 2. Some definitions for case classification were changed: 1) a contact case has onset of illness 4–75 days (formerly 4–60 days) after OPV was fed to a recipient in contact with the case; 2) "immunologically abnormal" includes cases in persons with proven or presumed immunological abnormalities regardless of the etiology of their abnormality (primary immunodeficiency, drug-induced suppression, etc.); and 3) "imported" includes cases in persons (US residents or others) with onset within 30 days before or after entry into the United States.

Comparison between ECPPC and ELCPPC Classification Systems

A total of 66 reported cases of paralytic illness with onset during the period 1980–86 met the CDC definition for paralytic poliomyelitis; 45 cases were confirmed by expert consultants in 1985 and the remaining cases by expert reviews since then. Poliovirus isolates were available for analysis for 46 (70 per cent) of the 66 cases, and van Wezel type-specific neutralization, oligonucleotide fingerprinting, and/or genomic sequencing was done. Forty-three viruses (93 per cent) were characterized as vaccine-like and three (7 per cent) as wild. These cases were classified according to the new ELCPPC system (Table 2).

The cases also were classified according to the old system, ECPCC, using only the available epidemiologic information (Table 1). Six cases were classified into the category "endemic, not vaccine-associated." In fact, poliovirus was isolated and characterized as vaccine-like from five of these cases. No specimen was available from the sixth patient. Three of the five cases classified as imported cases had poliovirus isolated and characterized as wild. No specimens were available from the other two patients. Five of the eight immune-deficient cases had received oral polio vaccine (OPV) and two had contact with a recipient; information was insufficient on the remaining case. All eight immune-deficient cases had poliovirus isolates characterized as vaccine-like.

For the period 1980–86, 60 of the 66 reported cases of paralytic poliomyelitis in the United States were vaccine-associated either by epidemiologic and/or strain characterization; five (8 per cent) cases were imported and one case that occurred in 1980 could not be classified as vaccine-associated paralytic poliomyelitis because no specimen for isolation was available and epidemiologic information was incomplete. While the ELCPPC classified 91 per cent of the cases by epidemiological and/or laboratory characterization as vaccine-associated, only 71 per cent were classified as vaccine-associated by the ECPPC which relies on epidemiological information (Table 3).

International Application

The World Health Organization (WHO) estimates that more than 250,000 cases of paralytic poliomyelitis occur each year worldwide.[15] Control or elimination of the disease in developed countries using either oral or inactivated polio vaccine has generated interest in the global eradication of poliomyelitis.[16,17] Given the progress by the Expanded Program on Immunization of WHO in achieving immunization with three doses of polio vaccine of more than 50 per cent of children in many developing countries,[15] the World Health Assembly adopted the goal in May 1988 of global eradication of poliomyelitis by the year 2000.[18] The WHO Region of the Americas, through the Pan American Health Organization (PAHO), established a target for elimination of indigenous poliomyelitis by the end of 1990, while the WHO European and Western Pacific Regions have set targets for elimination by 1995. Global eradication of poliomyelitis is feasible by achievement and maintenance of high immunization levels, effective surveillance to detect all new cases, and a rapid vigorous response to the occurrence of new cases.[19] To monitor these efforts, laboratory systems have been developed that identify which type(s) of polioviruses are circulating in a country or region, and to distinguish vaccine-virus from wild-virus,[14] but these methods need to be made more widely available. Countries without current access to such a laboratory system will face delay in being able to fully utilize the ELCPPC. However, even in the absence of such laboratory support, paralytic polio cases can be classified within the distinct epidemiologic categories of the ELCPPC.

TABLE 3—Comparison of Vaccine-Associated Paralytic Poliomyelitis Cases by Classification Category in the ECPPC and ELCPPC Systems

Category	System	Number of cases	
		Total	Vaccine-Associated (%)
Endemic*	ECPPC	53	47 (89%)
Sporadic*	ELCPPC	53	52 (98%)
Immunodeficient#	ECPPC	8	0** (0%)
Immunologically Abnormal#	ELCPPC	8	8 (100%)
Epidemic - Imported	ECPPC	5	0 (0%)
	ELCPPC	5	0 (0%)
All Categories	ECPPC	66	47 (71%)
	ELCPPC	66	60 (91%)

#Comparable categories
*Comparable categories
**Classification not available under ECPPC System

*A patient must have had paralysis clinically and epidemiologically compatible with poliomyelitis and, at 60 days after onset of symptoms, had a residual neurological deficit, had died, or had no information available on neurologic residua (This classification was formerly known as the Best Available Paralytic Poliomyelitis Case Count (BAPPCC) case definition).

Conclusion

Advances in methods to characterize strains of poliovirus as vaccine-like or wild provided the stimulus to improve upon the ECPPC system which relies entirely on epidemiological information. In addition, the epidemiology of poliomyelitis in the United States changed substantially. Epidemic and endemic wild poliomyelitis disease no longer exists in the United States, and vaccine-associated disease is the most frequent form of paralytic poliomyelitis. Since 1980 an average of only eight cases have been reported annually. The new system, ELCPPC, corrects for misclassifications in the older ECPPC system in which some cases were categorized as "endemic, not vaccine-associated" even when laboratory confirmation of a vaccine-virus infection was available.

While the ELCPPC format is at first glance somewhat complex, the greater descriptive accuracy of each category and subcategory provides more specific and useful information particularly concerning vaccine-associated disease. Information on the relative occurrence of wild-virus and vaccine-associated cases within each epidemiologic category can be obtained rapidly from this tabular format.

As current worldwide efforts to immunize children against polio are expanded and the effectiveness of these efforts is documented by a sustained decrease in polio cases, the ELCPPC system will become useful as a standardized tool to track the success of these programs and eventually to help in certifying countries or regions as free of wild poliomyelitis.

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Public Health Briefs

The Relationship between Women's Attitudes about Condoms and Their Use: Implications for Condom Promotion Programs

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Abstract: A survey of 759 women attending contraceptive care clinics revealed that a majority of women endorsed condom use as an important way to reduce the spread of AIDS (acquired immunodeficiency syndrome); 82 per cent reported past use of condoms, but only 14 per cent reported using condoms in addition to another form of contraception to prevent infection with sexually transmitted pathogens, including HIV (human immunodeficiency virus). Acceptance of condom advertisement, perceived male and peer acceptance of condoms, and the effect of condom use on the enjoyment of sexual intercourse predicted condom use. (*Am J Public Health* 1989; 79: 499–501.)

Introduction

In a recent mailing to American citizens on the subject of AIDS (acquired immunodeficiency syndrome), readers were advised that "not having sex," or having sex with "one mutually faithful, uninfected partner" are "safe behaviors" that will protect one from AIDS.[1] For individuals falling outside of these categorizations, condom-protected intercourse is the next best option. However, merely warning persons about the dangers of sexually transmitted HIV (human immunodeficiency virus) infection and urging the use of condoms may not result in their widespread adoption. We conducted a survey of sexually active women to assess their knowledge of, attitudes about, and use of condoms.

Methods

During the months of September and October 1987, 800 questionnaires pertaining to condom use knowledge, attitudes, and practice were distributed to clients at eight contraceptive care clinics (100 consecutive clients per site) in urban, suburban, small town, and college campus locations in southwestern Pennsylvania. The questionnaire consisted of 10 true/false questions, 12 attitudinal items, questions about past and current condom use, and several questions relating to the demographic characteristics of the applicants.

Categorial variables were analyzed using the chi-square test. Multiple linear logistic regression, using maximum likelihood estimation, was performed to simultaneously assess the association of co-variables with the dependent outcome of current condom use.[2]

Results

Seven hundred fifty-nine completed surveys were returned for analysis. Over half of the participants were between the ages of 19 and 25, 84 per cent were White, 41 per cent had at least a high school diploma, 15 per cent were married, and 15 per cent were living with a male partner to whom they were not married. 86 per cent reported a single sexual partner, 6 per cent had more than one sexual partner, and 8 per cent were not sexually active at the time of the survey.

Eighty-two per cent of the women reported that they had used condoms with a male sexual partner at least once in the past; 74 per cent reported that this was for "birth control," 3 per cent reported it was "to prevent VD," and 23 per cent said "for both purposes." Twenty-one per cent of the women reported they were currently using condoms with or without contraceptive foam for "birth control" and 14 per cent were currently using condoms and another contraceptive "to prevent VD."

Although the results from the true/false questions revealed a relatively high general level of knowledge about condoms and HIV transmission, a number of inaccuracies were noted, including: 26 per cent believed that vaseline "is the best lubricant to use with condoms"; 21 per cent believed that withdrawing the penis before ejaculation "is as safe" as using a condom in terms of preventing AIDS; 12 per cent believed that the virus which causes AIDS cannot be found in both male and female genital secretions; and 9 per cent thought a woman "could not get AIDS" from a male sexual partner unless she had anal intercourse with him.

Table 1 outlines the women's responses to the attitudinal items as a composite and then dichotomized between "users" (women who report current use of condoms with or without foam as a contraceptive) and "non-users." The majority endorsed condoms as an important way to prevent the spread of AIDS, especially with a new sexual partner, although 37 per cent expressed uncertainty of their ability to initiate use, and 22 per cent believed they would be "too embarrassed" to purchase condoms in a drug store. Non-users were more likely to feel that most men do not like using condoms, that condom-protected intercourse would not be enjoyable, were less likely to report peer acceptance of condoms, and more likely to be unsure about feeling "insulted" if a partner wanted to use a condom.

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TABLE 1—Attitudes Pertaining to Condom Use: Per Cent of Total Cohort and Condom Users versus Non-users

Women's Responses		% Agree	% Disagree	% Not sure
Most men do not like using condoms		80.5	7.4	12.1
	User Non	73.5 82.3	13.2 5.8	13.2 11.8
I would be insulted if my male partner wanted to use a condom while he was having sex with me		5.2	86.1	8.7
	User Non	6.0 5.0	90.7 84.9	3.3 10.1
If I were planning on having sex with a man I didn't know, I would ask him to use a condom		87	1.4	11.6
	User Non	91.4 85.9	7.1 6	7.9 12.6
I wouldn't know how to bring up the subject of condoms with a new sexual partner		14.9	63.2	21.9
	User Non	14.6 14.9	65.6 62.6	19.9 22.5
Most of the women I know think that using condoms with a new sexual partner is a good idea		77.2	6.1	16.6
	User Non	84.8 75.3	5.3 6.3	9.9 18.4
I would be too embarrassed to buy condoms at a drug store		21.8	62.8	15.4
	User Non	20.5 22.1	62.9 62.8	16.6 15.1
I worry a lot about getting AIDS from a new sexual partner		69.9	17.7	12.3
	User Non	69.8 70.0	18.8 17.5	11.4 12.6
I think that condoms should be advertised on television as a way to prevent the transmission of AIDS		84.0	5.2	10.8
	User Non	88.7 82.8	1.3 6.2	10.0 11.0
I think I would feel "dirty" if a man wanted to use condoms with me		2.5	93.5	4.1
	User Non	1.3 2.7	95.4 93.0	3.3 4.3
I think that condoms are an important way to reduce the spread of AIDS		95.6	1.2	3.1
	User Non	96.0 95.5	7.1 4	3.3 3.1
I don't think I would enjoy intercourse if my partner used condoms		12.7	70.3	17.0
	User Non	4.6 14.8	88.7 65.5	6.6 19.7
Even if I were using some other method of birth control, I think that I would still use a condom with a new sexual partner to prevent myself from becoming infected with VD (including AIDS)		81.2	3.7	15.1
	User Non	84.0 80.5	3.3 3.8	12.7 15.7

Differences by race in response to the knowledge questions were negligible (data available upon request to authors).

Responses to nine of the 12 attitudinal items differed by race (data available upon request to authors). Black respondents (n=117) were more likely to worry about AIDS, to endorse the use of condoms, especially with a new sexual partner, to report peer acceptance of condoms, and were less likely to feel unsure about initiating condom use or to be embarrassed about purchasing condoms.

Forward stepwise multiple logistic regressions were used to investigate the association between the previously described variables and condom use (Table 2). The first model contained demographic, knowledge, and attitudinal variables; the second contained only attitudinal variables. In both instances, attitudes alone were predictive of condom use.

Discussion

Among men, a variety of reasons have been recorded for underutilization of condoms, including: lack of ready access; failure to plan for sexual intercourse; the use of alcohol and/or drugs prior to sex; beliefs that condoms impair sexual pleasure; the belief that condoms would be offensive to one's partner; and the belief that condoms are unnatural or immoral.[5-7] Many of these same barriers no doubt hold true for women as well.[8]

Our finding that attitudes about condoms were the only independent variables capable of predicting their use suggests that condom promotion programs targeted to women must attempt to alter these attitudes. To that end, programs

TABLE 2—Variables Independently Associated with Condom Use: Results of Multiple Logistic Regression

	Odds ratio	95% CI
A. Terms examined: Age; Race; Educational Level; Location of Clinic; Marital Status; Number of Correct True/False Questions; Perceived Male Acceptance of Condoms; Intention to Use Condoms with New Male Partners; Peer Acceptance of Condoms; Barriers to Purchase; Condoms as Important Means of Preventing AIDS		
Male Acceptance[a]		
Agree	1.00	
Disagree	2.45	1.34, 4.49
Not Sure	1.19	0.67, 2.09
TV Advertisement[b]		
Agree	1.00	
Disagree	0.11	01, 0.79
Not Sure	0.88	0.48, 1.61
B. Terms Examined: Twelve Attitudinal Factors Only (Table 1)		
Enjoyment[c]		
Agree	1.00	
Disagree	4.26	1.92, 9.48
Not Sure	1.11	0.41, 3.06
Peer Acceptance[d]		
Agree	1.00	
Disagree	0.67	0.29, 1.57
Not Sure	2.14	1.17, 3.90

a) "Most men do not like using condoms."
b) "I think that condoms should be advertised on TV as a way to prevent the transmission of AIDS."
c) "I don't think I would enjoy intercourse if my partner used condoms."
d) "Most of the women I know think that using condoms with a new sexual partner is a good idea."

should include skills-training sessions which incorporate role playing and small group discussion. Experienced peers may be especially effective in addressing the perception that condoms interfere with sexual pleasure, since condom acceptance is known to increase with increased experience.[9]

Condom promotion programs must also be flexible enough to meet individual needs. There may be significantly different perceptions about condom use among women from different racial/ethnic backgrounds.

The observed discrepancy between the widespread endorsement of condoms and the frequency of reported use probably reflects the fact that many of the women we surveyed do not perceive themselves to be at risk for HIV infection,[10] because most reported only a single sexual partner. Therefore, the ability to generalize these findings to women who belong to targeted "risk groups" is limited. Nevertheless, we believe that the identification of generic barriers to women's use of condoms for contraception does provide direction for the development of condom promotion programs which can then be applied to other populations, especially those at increased risk of HIV infection.

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Sexual Behavior Changes and HIV Antibody in a Cohort of New York City Gay Men

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Abstract: A sample of 357 gay/bisexual men, ages 20 to 65, were interviewed in 1985 regarding their sexual patterns over the course of the AIDS (acquired immunodeficiency syndrome) epidemic. Within four months after the interview, men were tested for antibody to HIV (human immunodeficiency virus) and 36 per cent were found positive. Cessation of receptive anal intercourse was associated with HIV negative antibody status, (OR > 3.0). Reducing the number of sexual partners, becoming monogamous, or eliminating other forms of sexual contact were not related to risk of HIV. (*Am J Public Health* 1989; 79:501–503.)

Introduction

Without a vaccine against human immunodeficiency virus (HIV) or efficacious treatments for HIV-related illnesses, the primary means available for controlling the AIDS (acquired immunodeficiency syndrome) epidemic is through education and the self-regulation of transmission risk behavior. For gay/bisexual men, two main types of behavior change have been recommended[1–3]: reduce the number of different sexual partners, and eliminate the exchange of body fluids during sexual encounters. Although gay men in San Francisco and New York City have substantially changed their sexual behavior patterns to conform with these recommendations,[4–7] persistence of high-risk sexual activity continues among a minority of these well informed cohorts.[6] In addition, high rates of risk-taking behavior have been reported among gay men in areas with low AIDS incidence.[8,9]

While it is well established that unprotected anal intercourse is the primary sexual behavior associated with HIV infection,[10–15] evaluation of the usefulness of actual *changes* in sexual behavior in preventing HIV infection among those initially at high risk is less common.[16] Thus, we undertook this study to determine the extent to which changes in sexual behavior occurring after the onset of the AIDS epidemic altered the likelihood of HIV infection among gay men.

Methods

This analysis is based on 357 of the 746 gay men recruited to study the impact of AIDS on the New York City gay community.[7] Respondents were primarily White (87 per cent), college educated (81 per cent), with an average age of 36 (SD = 7.4). No differences were found between the groups who agreed to and who refused testing (but agreed to be interviewed) on any sexual behavior variable, including numbers of partners or frequency of anal intercourse. (See Martin, *et al.*[17] for a detailed discussion of this sample.)

HIV Antibody Serology

HIV testing was performed by the New York City Department of Health. Two HIV-ELISA (enzyme-linked

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A Comparison of Acquired Immunodeficiency Syndrome and Kaposi's Sarcoma Incidence Rates, Atlanta, 1983–86

Wong-Ho Chow, PhD, Jonathan M. Liff, PhD, Raymond S. Greenberg, MD, PhD, and Brian O. Williams, MD

Abstract: Recent temporal trends in incidence rates for acquired immunodeficiency syndrome (AIDS) and Kaposi's sarcoma were compared in metropolitan Atlanta, Georgia. From 1983 through 1986 the age-adjusted incidence rate of Kaposi's sarcoma among White males ages 25–49 increased 11-fold (2.2–24.4/100,000 person years). The annual incidence rate of AIDS increased ten-fold (11.3–113.3/100,000). These findings do not support earlier reports that the proportion of AIDS patients with Kaposi's sarcoma has decreased over time. (*Am J Public Health* 1989; 79:503–505.)

Introduction

In 1981, the first cases of acquired immunodeficiency syndrome (AIDS) in the United States were recognized among young homosexual males with Kaposi's sarcoma or *Pneumocystis carinii* pneumonia and no other known cause of immunodeficiency.[1,2] The proportion of reported AIDS cases with Kaposi's sarcoma has declined over time in the United States.[3,4] It has been suggested that some portion of this decreasing occurrence of AIDS-related Kaposi's sarcoma may be due to reduced exposure to cofactors of Kaposi's sarcoma,[4,5] such as the use of butyl nitrite among homosexual males.[6]

Few estimates of the incidence of AIDS have been reported because of difficulties in determining the size of the populations at risk.[2,6] Estimates of temporal trends in the occurrence of AIDS usually are based solely on changes in the total number of reported cases.[2,7,8] In contrast, trends over time in the incidence of Kaposi's sarcoma have been reported for the years prior to the onset of the current AIDS epidemic.[9] In the present study, recent temporal trends in incidence rates for both AIDS and Kaposi's sarcoma are reported and compared for young White males in metropolitan Atlanta, an area that has a relatively high incidence of AIDS.[10]

Methods

The data on AIDS incidence presented in this report were collected by the AIDS Surveillance Unit of the Georgia Department of Human Resources. AIDS became a reportable disease by law in Georgia in January 1984. For each case, the diagnosing physician is required to complete a standard reporting form with demographic and clinical information. AIDS cases diagnosed prior to 1984 were identified retrospectively by active surveillance methods including: review of tumor registry reports for Kaposi's sarcoma cases under 60 years of age at diagnosis; survey of physicians by letter and telephone; review of death certificates from Georgia; and review of requests received by the Centers for Disease Control (CDC) for pentamidine isethionate, a medication used for the treatment of *P. carinii* pneumonia[1,2,11] for which CDC has been the sole supplier until recently.[11]

Since first published in September 1982,[12] the AIDS definition for surveillance purposes has undergone minor revisions. Because the current study included cases diagnosed only through 1986, the latest revision would have had minimal impact on case identification.

The data on Kaposi's sarcoma presented in this report were collected by the Georgia Center for Cancer Statistics, a population-based cancer registry affiliated with the Surveillance, Epidemiology and End Results (SEER) Program of the National Cancer Institute.[13] This registry has been responsible for obtaining information on all cancers diagnosed among residents of metropolitan Atlanta since 1975.[14]

The primary method of cancer patient identification was through review of pathology reports and discharge diagnoses at the hospitals, clinics, and freestanding pathology laboratories in and around metropolitan Atlanta. Overall, less than 1 per cent of cases reported to the registry were identified from death certificates only. The abstracted demographic and clinical information was transcribed onto standard data collection forms, edited, and coded according to the guidelines of the SEER Program.[13]

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TABLE 1—Demographic Characteristics of AIDS and Kaposi's Sarcoma Patients Diagnosed between 1983 and 1986, Metropolitan Atlanta, Georgia

Characteristics	Category	All AIDS n	All AIDS (%)*	Kaposi's Sarcoma n	Kaposi's Sarcoma (%)*
Race	White	377	74.8	74	85.1
	Black	119	23.6	10	11.5
	Other**	8	1.6	3	3.4
Sex	Male	490	97.2	86	98.9
	Female	14	2.8	1	1.1
Age at Diagnosis (years)	<25	30	6.0	3	3.2
	25–29	102	20.2	21	22.3
	30–34	137	27.2	20	21.3
	35–39	127	25.2	15	16.0
	40–44	61	12.1	17	18.1
	45–49	28	5.6	8	8.5
	50–54	10	2.0	1	1.1
	55–59	9	1.8	2	2.1
County of Residence	Fulton	364	72.2	67	77.0
	DeKalb	89	17.7	14	16.1
	Other†	51	10.1	6	6.9

*Numbers in parentheses are percentages of all patients within a disease classification
**Includes 2 Kaposi's sarcoma cases with unknown race; all AIDS cases under this category were Hispanic
†Includes 4 AIDS cases with unknown county of residence

For the current analyses, information about all histologically confirmed Kaposi's sarcomas newly diagnosed between January 1, 1983 and December 31, 1986 among persons under 60 years of age was extracted from the database of the cancer registry. Similarly, information about all residents under 60 years of age with a new diagnosis of AIDS during the same time period was extracted from the database of the AIDS Surveillance Unit. No linkage of information about individuals reported to the cancer registry and the AIDS Surveillance Unit was attempted because of confidentiality requirements for each database. Race and age-specific incidence rates of Kaposi's sarcoma and AIDS were computed separately, using as denominators estimates derived by the US Bureau of the Census and the National Cancer Institute for the years 1983–86. The incidence rates for each calendar year were age-adjusted to the 1970 US population by the direct method, and approximate 95 per cent confidence intervals were calculated.[15]

Results

This study population included 504 patients with AIDS and 87 patients with Kaposi's sarcoma. A slightly higher proportion of Kaposi's sarcoma patients were White (85.1 per cent) compared with all AIDS patients (74.8 per cent). The distributions by sex, age, and county of residence were similar for these two samples of patients, however (Table 1). The median age for both groups of patients was 34 years, with the age ranging from 1 to 59 years for AIDS cases and from 19 to 58 years for Kaposi's sarcoma cases.

The remaining analyses focused only on White males between the ages of 25 and 49 years, who lived in Fulton or DeKalb counties; among them, the peak incidence rate of AIDS during 1983–86 (92.5 per 100,000) occurred among those ages 35–39, while that of Kaposi's sarcoma (19.2 per 100,000) peaked at ages 40–44 years.

Between 1983 and 1986, the age-adjusted incidence rate of Kaposi's sarcoma increased 11-fold, from 2.2 to 24.4 cases per 100,000 person years (Table 2), and there was a 10-fold rise in the annual incidence rate of AIDS, from 11.3 to 113.3. The incidence rate ratios of Kaposi's sarcoma and AIDS varied from 0.19 to 0.29, with no consistent trend over time (Table 2).

Discussion

In contrast to previous reports,[3,4] we found parallel increases in the incidence of AIDS and Kaposi's sarcoma among White males ages 25–49 years in Atlanta. If AIDS cases had been underreported in metropolitan Atlanta to a greater extent than in other cities, and if the underreporting had increased each year, the parallel increase we observed could have occurred even if the true ratio of Kaposi's sarcoma to AIDS cases had been decreasing. Data from CDC,[16] however, show that the proportional increase in the number of AIDS cases in Atlanta between 1984 and 1986 was even greater than in the nation as a whole. In addition, there was no abrupt increase in the incidence of reported AIDS cases in 1984, the year during which AIDS became reportable by law in Georgia, suggesting that the reporting of AIDS in Atlanta was fairly complete even prior to 1984. Previous studies in other settings also have found that the reporting of AIDS is quite complete in general.[17,18]

As conditions other than Kaposi's sarcoma and *P. carinii* pneumonia have been accepted as diagnostic evidence of AIDS, the proportion of cases presenting with Kaposi's sarcoma might be expected to decrease over time; we found that this proportion decreased from 26.3 per cent in 1983 to 19.0 per cent in 1986 among White males. Our data, based on separate data sources include Kaposi's sarcomas presenting subsequent to the initial diagnoses of AIDS, thus more clearly indicating the proportion of AIDS patients with

TABLE 2—Annual Numbers of Cases, Age-Adjusted Incidence Rates, and Rate Ratios of Kaposi's Sarcoma and AIDS among White Males Ages 25–49 in Fulton and DeKalb Counties, Metropolitan Atlanta, 1983–86

Group		Year of Diagnosis 1983	1984	1985	1986
Kaposi's Sarcoma	Number of cases	3	12	20	31
	Age-adjusted incidence rate*	2.2	8.5	16.7	24.4
	(95% CI)	(0.0, 4.7)	(3.6, 13.3)	(9.1, 24.4)	(15.5, 33.3)
AIDS	Number of Cases	16	42	105	148
	Age-adjusted incidence rate*	11.3	32.4	79.3	113.3
	(95% CI)	(5.7, 16.8)	(22.3, 42.4)	(63.7, 94.9)	(94.5, 132.1)
	Ratio of Kaposi's sarcoma: AIDS	0.19	0.29	0.19	0.21
	(95% CI)	(0.06, 0.56)	(0.16, 0.52)	(0.12, 0.29)	(0.15, 0.30)

*per 100,000 person years

Kaposi's sarcoma. Since the trend analysis was restricted to White men ages 25–49, it is highly likely that the registered patients with Kaposi's sarcoma also had AIDS.

When cases from other demographic subgroups (i.e., females and Blacks) were included in the analysis, the ratio of Kaposi's sarcoma to AIDS decreased somewhat in the later two years, from 0.19 in 1983 to 0.16 in 1986. The occurrence of Kaposi's sarcoma and AIDS in these other demographic subgroups, however, was too small and too recent to permit the identification of any meaningful associations.

The findings of this study suggest that the proportion of AIDS patients with concomitant Kaposi's sarcoma in young White males has not changed over time. Since these results differ from those in earlier reports, further analyses in other populations are warranted.

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The Hawaii Chlamydia Network Project: A Successful Program Incorporating Close Intra-agency Cooperation

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Abstract: The Hawaii State Department of Health's Chlamydia Network Project screened 272 asymptomatic females for chlamydia; 20 (7.4 per cent) were found to be positive. When interviewed, 75 per cent (15/20) of the cases gave information such that their sexual partner(s) could be located. Ninety per cent (14/16) of the locatable partners were brought to examination within seven days. Keys to success were the training of family planning clinic staff in STD (sexually transmitted disease) control methods, and close intra-agency cooperation. (*Am J Public Health* 1989; 79:505–507.)

Introduction

In October 1986, the Hawaii State Department of Health began a Chlamydia Network Project. The purposes of this project were to:

- Ascertain the prevalence of clinically inapparent chlamydial infections in female family planning patients;
- Develop screening protocols for chlamydia in all state-subsidized family planning clinics; and
- Develop a "network" between the Office of Family Planning and the Sexually Transmitted Disease (STD) Control Program in order to ensure all screen positive female clients and their locatable sex partners are brought to treatment in a timely manner.

Methods

The five major state-subsidized family planning clinics on the island of Oahu were involved. A total of 272 asymptomatic women were screened by obtaining a weighted sample from each of the participating clinics between December 1986 and April 1987. Eligibility criteria included any female having one or more of the following high-risk factors:[1–9] less than 25 years old; single; more than one sexual partner in the preceding 60 days; past history of a sexually transmitted disease (STD); patient being seen for an STD or STD screen, but who otherwise would not be treated for chlamydia.

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Tuberculosis Surveillance in a State Prison System

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Abstract: After four inmates at two New Mexico prisons converted their tuberculosis skin tests, a mass screening program was carried out at all of the State's adult correctional facilities (2,240 inmates). Previously unknown converters were found with a disproportionally high per cent of converters (6.8 per cent and 6.3 per cent) and reactors (14.4 per cent and 12.2 per cent) at the first two prisons. No index cases were found. State policy has been revised to include two-stage skin testing of new inmates, annual testing thereafter, and screening of all new correctional staff. (*Am J Public Health* 1989; 79: 507–509.)

Introduction

Although tuberculosis has greatly diminished in frequency in the general population of this country in recent decades, it remains a significant problem in certain population groups: some minorities, immigrants from third world countries, patients with human immunodeficiency virus (HIV) infection, residents of nursing homes, and other institutionalized groups.[1–5] Several authors have called attention to the problem of contagious tuberculosis in jails and prisons and the need for surveillance programs.[6–10] Stead in the late 1970s[6,7] showed that tuberculosis morbidity was 6.5 times greater in Arkansas prisons than in the general population of that state.[7]

The State of New Mexico Corrections Department operates seven adult correctional facilities. Prior to August of 1986, tuberculosis screening of new inmates consisted of a single Mantoux test (5 TU, PPD—Tuberculin). Those with a reaction of 10 mm of induration or more and those who gave a history of a positive skin test were x-rayed, after which a decision was made about further investigation and prophylactic treatment.[11,12] Those with negative reactions were scheduled to have repeat skin testing every two years. The tuberculin positivity rate in 1985 was 10.5 per cent.

In September of 1986, four cases of skin test conversion while in prison came to our attention. These four men all had negative skin tests at the same prison (SNM) in 1984, and one of them had been transferred shortly thereafter to another prison in the system.

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We decided to conduct a mass screening program at SNM facility. The results obtained led us to extend the program to the other five men's facilities and to the women's prison. The screening was carried out between October 1986 and April 1987.

Methods

Medical records of all inmates were reviewed; those who had been skin tested with a 0 millimeter result in the previous three months were not retested, but are included in the data for the screening program results. Inmates with a documented reaction of 10 mm or more in the past and with a chronic cough had a chest x-ray done and three sputum specimens obtained for smear and culture. Documented reactors without a cough who had completed a full course of INH (isoniazid) received no further investigation. Reactors who had not completed a course of INH and had not had a chest x-ray in the past six months, were x-rayed at the time of the mass screening. All inmates with a documented negative PPD (purified protein derivative of tuberculin) reading prior to three months earlier were skin tested. Those with any palpable reaction of less than 10 millimeters had a second stage test applied at least one week after the first. Inmates who had not been tested at intake because they gave a history of a positive skin test were tested. All converters and newly confirmed reactors also received a chest x-ray, and sputum investigation if the x-ray was suspicious. A converter was defined as one with a significant reaction whose previous negative test was within the past two years; a reactor as one whose reaction was of more than two years or of unknown duration. The housing location of all reactors and converters was recorded.

An effort was made to screen all employees at all seven correctional facilities. Those employees who could document or give a reliable history of a previous positive skin test were not retested, but did have a chest x-ray. All others received two-stage skin testing. Chest x-rays were also ordered on all reactors and converters.

Results

Inmate Screening

Table 1 shows the distribution of reactors and converters among the inmates of the seven adult correctional facilities in the state. All inmates participated in the screening program.

The skin test positivity rate (reactors plus converters divided by total number tested) was 10.7 per cent or lower in all but two of the facilities; one prison (SNM) was the one in which converters had first been identified, and the other (CNM) was the prison to which one of the converters had been transferred.

TABLE 1—Results of Inmate Screening

Facility	Screened	Negative	Reactors		Converters		Positivity Rate (%)*
			No.	(%)	No.	(%)	
SNM	466	367	67	(14.4)	32	(6.8)	(21.2)
CNM	574	468	70	(12.2)	36	(6.3)	(18.5)
PNM	685	612	62	(9.1)	11	(1.6)	(10.7)
CCW	142	129	9	(6.3)	4	(2.9)	(9.2)
CSB	61	56	5	(8.2)	0	(0)	(8.2)
LLCC	230	214	15	(6.5)	1	(0.5)	(7.0)
RCC	82	77	2	(2.4)	3	(3.7)	(6.1)
TOTALS	2,240	1,923	230	(10.3)	87	(3.9)	(14.2)

*Positivity Rate = (No. Reactors plus No. Converters) divided by No. Screened.
SNM, CNM, and PNM are medium or maximum security facilities for men. CCW is the Correctional Center for Women, all custody levels. CSB, LLCC, and RCC are minimum security facilities for men.

No active cases were found as a result of the chest x-rays on 269 reactors and converters, and sputum examinations on 38. The converters were scattered randomly among housing units or cell blocks, giving no clues to possible location of an index case.

Forty-nine (2 per cent) of the 2,240 screened inmates had no record of a previous skin test; 118 (5 per cent had a last negative skin test over two and a half years previously. (Since data were incomplete, the true number may have been higher.) Twenty of these 118 had significant reactions in the screening program.

Employee Screening

Among the employees the level of compliance was high (Table 2) and the skin test positivity rate ranged from 6.7 to 14.5 per cent, with an overall rate of 10.5 per cent. Most of the employees had not been skin tested for many years, if ever, and were therefore classified as reactors if they had a significant reaction in the screening program. Chest x-rays were done on 120 employees. No active cases of tuberculosis were found.

Discussion

Although we had thought that there might be undiscovered contagious cases in the prison system, there are other possible explanations for the skin test conversions. These include the use of inactive or out-dated antigen for the initial test, improper placement (subcutaneous rather than intradermal), and the booster effect of the initial test.[13–16] These factors do not explain the discrepancy between the results at CNM and SNM and the other five facilities, however. All of the initial testing was done at the same place, and the subsequent assignment to one of the seven facilities was not based on any factors which could conceivably influence converter rates. The possibility that over-reading of skin test reactions occurred at CNM and SNM seems unlikely. The education and preparation for this mass screening program was intensive and statewide. Most of the significant reactions were confirmed by a second observer, and most of the converters had induration of 15 to 25 mm, not borderline reactions.

Our failure to find any clustering of converters in individual housing units or any active index case or cases at CNM and SNM, while frustrating, was not too surprising. While prisons are close communities where contagious disease can spread easily, the turnover rate is high, and transfers from one facility to another within the system are frequent. The turnover rate among correctional staff is also high.

We concluded that the established tuberculosis screening policy had been inadequate. A few inmates had been missed in the intake testing program and more had been missed in the biennial retesting. Doubtful and insignificant reactions had not been retested at intake.

As a result of this experience, the tuberculosis surveillance program of the New Mexico Corrections Department has been revised as follows:

- Two-stage testing is done at intake on all inmates unless there is documentation of a past significant reaction. The first test is read at one week, and the second applied at that time to all with reactions of 0 to 9 mm.
- Inmates who give an undocumented history of a past positive test are retested at intake.
- Those with insignificant reactions in the two-stage initial testing are retested annually, with a single Mantoux test read at 48-72 hours. We chose annual testing because we want to monitor the current situation closely, and because we anticipate an increase in HIV infection. We want to be prepared to deal with the eventuality of reactivation tuberculosis in patients with HIV infection and to detect spread of tuberculosis in a timely fashion.
- Follow-up chest radiographs are done at one and two years on all those who refuse recommended INH, who are non-compliant with prophylactic treatment, or who have the drug discontinued for medical reasons. We are also intensifying our efforts to improve treatment compliance.
- A policy of screening all new correctional staff has been instituted, with two-stage skin testing.

TABLE 2—Results of Employee Screening

Facility	Employees	Screened		Negative	Reactors		Converters		Positivity Rate (%)*
		No.	(%)		No.	(%)	No.	(%)	
SNM	280	258	(92)	236	19	(7.4)	3	(1.1)	(8.5)
CNM	338	245	(72)	224	15	(6.1)	6	(2.5)	(8.6)
PNM	449	471	(94)	417	54	(11.5)	0	(0)	(11.5)
CCW & RDC	211	200	(95)	171	29	(14.5)	0	(0)	(14.5)
CSB	28	28	(100)	25	3	(10.7)	0	(0)	(10.7)
LLCC	75	75	(100)	70	4	(5.3)	1	(1.4)	(6.7)
RCC	46	46	(100)	41	5	(10.9)	0	(0)	(10.9)
TOTALS	1,477	1,323	(90)	1,184	129	(9.8)	10	(0.7)	(10.5)

*Positivity Rate = (No. Reactors plus No. Converters) divided by No. Screened.
SNM, CNM, and PNM are medium or maximum security facilities for men. CCW is the Correctional Center for Women, all custody levels. RDC is the Reception and Diagnostic Center for all new arrivals, men and women. CSB, LLCC, and RCC are minimum security facilities for men.

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Immunization Coverage and Social Differentiation in Urban Senegal

DIDIER FASSIN AND EMILE JEANNEE

Abstract: We studied sociocultural factors associated with immunization coverage among a random sample of 500 mothers in the underprivileged suburbs of Dakar, Senegal. Factors associated with immunization of children under age 5 were maternal educational level and socioeconomic conditions (especially husband's regular salary); maternal age on arrival in town was more influential than total number of resident years in town; no effect of ethnicity was evident. (*Am J Public Health* 1989; 79:509–511.)

Introduction

Health behavior is strongly influenced by social variables, hence could also be used as an indicator of social differentiation. In developing countries, the recent growth of urbanization has created new social structures to which differences of origins, ethnicity, cultural background, economic level, age at migration, and duration of residence in town contribute in varying degrees. Classical studies in urban anthropology have shown how urbanization modifies traditional society and promotes different social practices,[1–3] but most studies of health behavior have taken little account of these changes,[4–6] as a recent review of literature on health interviews surveys showed;[7] nevertheless, comparative data taking socio-cultural variables into account are necessary for identification of risk factors and risk groups.

Immunization coverage is an objective indicator of health behavior in urban Senegal because differences in coverage cannot be related only to financial considerations (vaccines cost little) or to physical access (the WHO Expanded Program on Immunization covers the whole town).

The study to be described was carried out as a part of a multidisciplinary approach to urbanization and health in Africa. Field research was completed in Pikine, an underprivileged suburb of Dakar where slums of the capital were rehoused in the 1960s and 1970s. Population was estimated at 600,000 inhabitants in 1985 with a doubling period of five years over the past two decades. Primary health care is available through a Belgium-Senegalese project; a national program of immunization has been in effect for many years.

Methods

Study Population

Five hundred mothers were randomly selected through two-stage cluster sampling. Because no census list existed, we used a list of the 260 administrative districts, selecting 25 at random. In each district drawn, two blocks were randomly chosen on an updated map of Pikine. In the 50 blocks thus selected, interviews were held with the first 10 mothers who had at least one child under 5 years old, starting at the eastern corner of the block. If such a mother was not at home after two visits, another woman from the same block was included. When a man had several wives, each of them was seen separately. Interviews were conducted jointly by two investigators, one of whom was a woman who lived in Pikine. Interview schedules included questions on sociocultural items and access to health care; in particular, the mothers were asked whether the youngest child between one and five

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years old had been immunized and the response was verified with the vaccination card in her possession.

Variables and Analysis

Sociocultural variables included ethnic group, age, marital status, maternal educational level, socioeconomic level, experience in town, and indicators of urban integration and village attachment. For *ethnic groups*, Wolofs and Lebus, Tukulors and Fulas, Bambaras and Soninkes, Diolas and Manjaks were respectively grouped together since the social structure and historical background of each member of the pair were closely related. For *socioeconomic level*, three different criteria were used: husband's regular salary, ownership of the house, possession of a couch; for *experience in town*, mother's age on first coming to town, the length of stay in town, the presence of child's father and or mother in town; for *urban integration*, use of Wolof as a communication language, participation in urban associations, knowledge of political local leader's name; for *village attachment*, use of mother tongue (for non-Wolof), participation in village associations, return to village in the past five years. Immunization coverage was considered absent if no vaccine had been given and present if partial or complete protection was noted against diphtheria, tetanus, whooping cough, poliomyelitis, measles, and yellow fever.

Results

Five hundred women were interviewed: eight were absent twice, 41 refused to answer, and thus 49 others were selected and interviewed. Sociocultural variables related to immunization coverage are: mother's educational level, husband's regular salary, ownership of their house, possession of a couch, mother's age on first coming to town, maternal parents living in town, mother-tongue speaking among non-Wolof, participation in village associations; recent return to village is negatively correlated to vaccination (Table 1). There are no important differences in immunization coverage among the largest ethnic groups (Table 2).

Discussion

Socioeconomic conditions and educational level are the main variables associated with immunization coverage. Although statistically correlated, the three criteria used for measuring the standard of living are not equivalent: having a salary means getting a regular income, as well as being protected by a social insurance for health problems; owning a house is less discriminatory, since middle-class households often prefer to rent an apartment in better districts, and poorer ones are very attached to any kind of property; finally, possessing a couch does not merely indicate objective "wealth" but also represents a status symbol.[8] The influence of socioeconomic conditions on immunization is not easy to understand at first analysis since vaccines only cost 100 FCFA, i.e., $0.3. Nevertheless, two strictly financial reasons can be proposed: although low in cost, this price may still prevent the poorest women from bringing their child to the dispensary for immunization; moreover, vaccines are sometimes unavailable through public outlets and people have to buy them in the private pharmacies where they cost 10 times as much. But poverty is also closely related to lack of education as is evident from this study; lack of education may prevent information about vaccination programs from reaching parents, as has been shown in Cameroon.[9]

Apparently the mother's early experience in town is a more important factor for her child's immunization than duration of her residence in town. The indicators of urban integration which we selected do not seem to have any influence on immunization coverage, perhaps because they were not discriminating enough; however, these same factors were associated with access to curative care for mothers and children in the same study.[10] They may be related to the ability to find one's way to a dispensary when ill, rather than attend a vaccination session whose existence is announced by loud-speakers. Our indicators of village attachment were supposed to measure the relationship of the women with their original rural community; but they apparently reflect different influences: recent return to village was negatively correlated to vaccination while the opposite was found for participation in a village association. The latter may indicate involvement in urban networks as well as traditional ties.[11]

Ethnicity is a central concept in social anthropology and traditionally considered as prominent in determinism of

TABLE 1—Immunization Coverage and Sociocultural Variables of Mothers in Urban Senegal

Socio-Cultural Variables	n	% of Children Not Vaccinated	Confidence Interval of the Difference
Mother's Age (years)			
<25	201	35.3	
≥25	291	34.7	(0.1, 1.2)
Marital Status			
monogamous	276	34.4	
polygamous	201	35.3	(0.1, 1.7)
Mother's Educational Level			
none	356	41.8	
at least primary	142	18.3	(19.8, 27.2)
Husband's Regular Salary			
no	317	42.6	
yes	180	21.7	(17.3, 24.5)
Owning of the House			
no	165	41.2	
yes	333	32.1	(6.6, 11.6)
Possession of a Couch			
no	328	40.8	
yes	170	24.1	(13.4, 20.0)
Mother's Age at First Arrival in Town			
after 10	244	40.6	
before 10	240	28.7	(9.0, 14.8)
Years Spent in Town			
less than 10	122	41.0	
more than 10	361	32.4	(6.1, 11.1)
Child's Parents Living in Town			
no	285	42.1	
yes	212	25.5	(13.3, 19.9)
Wolof-speaking (for non-Wolof)			
no	41	26.8	
yes	208	37.5	(6.9, 14.5)
Participation in Urban Associations			
no	396	35.9	
yes	102	32.3	(2.0, 5.2)
Knowledge of Political Local Leader's Name			
no	104	36.5	
yes	393	34.9	(0.5, 2.7)
Mother-tongue Speaking (for non-Wolof)			
no	39	20.5	
yes	208	38.0	(12.8, 22.2)
Participation in Village Associations			
no	455	37.6	
yes	43	9.3	(24.3, 32.3)
Return to Village in the past 5 Years			
no	205	26.8	
yes	285	41.0	(−11.1, −17.3)

N = 500

TABLE 2—Ethnic Groups, Immunization Coverage of Children, and Sociocultural Variables

Ethnic Group	N	% of Children Not Vaccinated	% Lower Socioeconomic Group	% No Educational Level
Wolof-Lebu	249	34 5	16 9	71 5
Serer	56	37 5	21 1	70 2
Tukulor-Fula	141	39 0	17 6	73 2
Bambara-Soninke	13	15 4	53 8	38 5
Diola-Manjak	26	23 1	11 5	53 9
Others	13	38 5	38 5	84 6
Total	498	35 1	18 8	70 4

behavior.[12] Nevertheless, in African cities, it has been shown that it is modified by other social factors: for example, unionism in the Cooperbelt area in South Africa is independent on the ethnic group of mine workers, but is correlated to the length of residence in an urban area;[13] an African townsman thus seems to be a townsman before being a member of his ethnic group.[14] In Senegal, Wolofs and Lebus represent the dominant group (43 per cent of total population) as well as the oldest one in the Dakar area. Tukulors and Fulas, on the other hand, have recently arrived in the capital from the Sahelian region. Serers are mainly peasants from the central part of Senegal. Other groups are minorities from Eastern (Soninkes, Bambaras) and Southern Senegal (Diolas, Manjaks). There is no difference in immunization coverage between Wolofs-Lebus and Tukulors-Fulas who have very similar socioeconomic levels, although their culture, history, and urbanization are quite different. And when all groups are considered in the same analysis, ethnicity does not appear to differentiate immunization coverage; the same findings were obtained for access to curative care.[10]

Determinants of health behavior in urban Africa are often described as "cultural." In the pluralistic medical system of cities, people would search for a therapy according to their ethnic and geographical origins. In the particular case of vaccinations which can be considered as an example of preventive health behavior, and in the specific context of the underpriviledged suburbs of Dakar, we have shown that educational level and socioeconomic conditions are correlated with immunization coverage, while ethnic distinctions are not a factor when determining risk groups in urban populations.

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Prevalence of Diabetes in a Navajo Indian Community

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Abstract: The age- and sex-adjusted prevalence of non-insulin-dependent diabetes mellitus (NIDDM) of 494 (76 per cent) Navajo adults living in a reservation community was 10.2 per cent, approximately 60 per cent greater than the estimated prevalence (6.4 per cent) in the general U.S. population. The screening protocol utilized likely underestimates the prevalence of NIDDM in this population. A high proportion of Navajo people were overweight when compared to the general U.S. population. (*Am J Public Health* 1989; 78:511–513.)

Introduction

Diabetes mellitus is a public health problem of increasing importance for Native Americans. Diabetes in Southwest Indian tribes is almost exclusively non-insulin-dependent diabetes mellitus (NIDDM).[1] Most data regarding the prevalence of NIDDM in Navajo Indians, the largest Indian tribe in the United States, have been based on hospital chart reviews or on limited population surveys.[2] We therefore conducted a community-based screening program for NIDDM in adults in a representative sample of Navajo Indians.

Methods

The study was carried out from April 1, 1986 to March 31, 1987 in the Teec Nos Pos (TNP) Chapter, Shiprock Service Unit, on the 25,000 square mile Navajo Indian Reservation in northeastern Arizona. The 360 square mile area is inhabited almost exclusively by Navajo Indians.

Study Subjects

The 1986 population of the TNP Chapter was estimated to be 1,540.* By extrapolation from the Shiprock Service

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Unit population, where an estimated 42 per cent of the population are 20 years of age or greater,** we estimated that 647 individuals in the TNP chapter were in this age group. The analyses were restricted to data for persons residing within the boundaries of the Chapter who were 20 to 74 years of age.

Approval to conduct the study was obtained from the local elected Health Board and the Navajo Area Indian Health Service institutional review board. Approximately 35 per cent of the subjects were screened at home and 13 per cent were screened at a local clinic. The remaining 52 per cent were screened in public places.

Screening was performed using capillary blood analyzed with a portable blood glucose monitoring system (Accu-Chek II Blood Glucose Monitors and Chemstrip bG Reagent Strips, Boehringer Mannheim, Indianapolis, IN). Fingersticks to obtain capillary blood were performed by nurses or other health care providers who were specially trained in the use of the instrument.

Individuals whose random blood glucose (RBG) levels were 7.8 mmol/l (140 mg/dl) or greater were referred to a clinical laboratory for fasting venous blood glucose (FBG) determination by a hexokinase method. Subjects whose FBG was greater than 7.8 mmol/l (140 mg/dl) were referred to Indian Health Service clinical providers for further evaluation and therapy.

The diagnosis of diabetes mellitus was made in persons with FBGs after an overnight fast of 7.8 mmol/l (140 mg/dl) or greater on two occasions, or with an RBG of 11.1 mmol/l (200 mg/dl) or greater plus classic signs and symptoms of diabetes mellitus.[3] Circumstances did not allow us to perform a formal glucose tolerance test. Blood glucose levels of subjects claiming to be diabetic at the time of screening were not always tested, although all diagnoses were subsequently confirmed by chart review according to the above criteria.

Weight and height were measured in subjects when possible. Weight-height index (W/H^p) was calculated in kg/m^p using $p = 2$ for males and $p = 1.5$ for females. Percentage of persons overweight at selected percentiles based on the weight-height index in the study population were compared to the United States rates.[4]

Age- and sex-adjusted prevalence rates were calculated using the 1980 US population as the reference population.[5] The US prevalence rates for diabetes were calculated for the above population using estimates of the numbers of diabetics aged 20–74 years.[6] Rates were age-adjusted to the 1980 US population by the direct method.

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Results

Screening was performed on 76 per cent (494 individuals) of the estimated population age 20 years or older. Follow-up glucose testing was performed on 94 per cent (15/16) of those with RBG of 10.0 mmol/l (180 mg/dl) or greater, and 86 per cent (44/51) of those with RBG of 7.8 mmol/l (140 mg/dl) or greater who were not previously known to be diabetic. (A 72 year old male with a RBG of 21.2 mmol/l (382 mg/dl) who refused any follow-up testing or clinical interview was not counted as a diabetic in the prevalence calculations.)

The age and sex distributions of those screened were similar to those estimated for the TNP population, except that the screened group had a disproportionately low number of subjects under age 30 years, and males were slightly underrepresented in the study population. The age- and sex-specific prevalence rates of diabetes mellitus in the Teec Nos Pos chapter and the general US population are shown in Table 1. Twelve new cases (7 women, 5 men) of diabetes were diagnosed, and 36 previously known diabetics (23 women, 13 men) were identified. The crude prevalence rate of NIDDM in adults age 20–74 years was 9.9 per cent. The age- and sex-adjusted prevalence rate of diabetes in the TNP population ages 20–74 years was 10.2 per cent compared to the 1980 US rate of 6.4 per cent.

Data adequate to compute a weight-height index were available for 236 females and 164 males. A higher percentage of individuals in both sexes and all age groups (except males ages 20–24) were overweight in the study population than in the US population (Table 2).

Discussion

While the burden of infectious diseases in American Indian communities has decreased substantially over the past three decades, chronic illnesses such as NIDDM have become increasingly prevalent.[7] Diabetes mellitus was thought to be almost nonexistent among Navajo Indians before the 1940s, but subsequent surveys have shown large increases in the disease prevalence.[2]

No post-load glucose tolerance tests were performed in the present study. In addition, the screening procedure did not require an FBG. Thus, the prevalence rates of diabetes in the TNP Chapter are not directly comparable to those estimated from National Health and Nutrition Examination

TABLE 1—Age-Specific Prevalence Rates of Diabetes Mellitus (per 100 population) in a Navajo Community, 1987, and in the 1976–1980 US Population

AGE (years)	20–74	20–44	45–54	55–64	65–74
US all races					
Both	6.6	2.0	8.5	12.8	17.7
Men	5.7	1.4	7.9	9.6	19.2
Women	7.4	2.5	9.0	15.5	16.5
Teec Nos Pos					
Both (%)	9.9	3.4	15.9	23.1	21.0
(N)	(44/444)	(7/254)	(13/82)	(15/65)	(9/43)
Men (%)	9.1	2.2	10	30.8	11.8
(N)	(16/176)	(2/93)	(4/40)	(8/26)	(2/17)
Women (%)	10.5	3.1	21.4	18.0	26.9
(N)	(28/268)	(5/161)	(9/42)	(7/39)	(7/26)

SOURCE: NHANES II

TABLE 2—Percent of Persons Overweight at Selected Percentiles, Teec Nos Pos Chapter, Compared to US All Races (NHANES I) by Sex and Age, Based on the Weight-Height Index*

Age (years)	TNP, >85 Percentile	US, >85 Percentile	TNP, >95 Percentile	US, >95 Percentile
Male				
20–24	9.1	11.5	0	3.4
25–34	36.4	21.8	21.2	6.6
35–44	41.2	28.2	20.6	5.1
45–54	64.9	26.9	29.7	6.8
55–64	39.1	23.8	8.7	7.4
65–74	53	21.9	6.7	4.4
20–74	42.1	22.8	17.1	5.8
Female				
20–24	23.7	12.7	5.3	3.8
25–34	50	20.8	15.5	7.0
35–44	58.2	30.3	14.5	10.5
45–54	69.4	35.8	27.8	9.1
55–64	60	41.3	26.7	10.5
65–74	84.2	39.9	31.6	9.1
20–74	54.7	29.5	18.2	8.3

*Weight-height in kg/m^p where p = 2 for males and p = 1.5 for females. Overweight criteria based on sex specific 85th and 95th percentile for persons age 20–29 years in NHANES I.[4]

Survey (NHANES II) data, which required that subjects meet the National Diabetes Data Group (NDDG) criteria after a 75 gram oral glucose load. Although our FBG 7.8 mmol/l (140 mg/dl) or greater is diagnostic of diabetes by NDDG diagnostic criteria, those criteria also allow the diagnosis of diabetes with a lower FBG plus elevated midtest and two-hour plasma glucose levels. Because only about one-fourth of the undiagnosed diabetics in the NHANES II study could be classified as diabetic solely on the basis of our FBG value, it is likely that the administration of a glucose tolerance test would have identified a higher prevalence of diabetes. In addition, because Type I (insulin-dependent) diabetes mellitus has rarely been described in full heritage southwestern American Indians,[1,8] the 1.6 ratio of the prevalence of NIDDM in the study population compared to that in the general US population, which includes persons with Type I diabetes, is underestimated.

The prevalence of NIDDM between the ages of 20 and 74 in the population of Navajo Indians studied exceeded that in the same age groups in the US population by approximately 60 per cent. The duration-specific prevalence rates of microvascular complications of NIDDM in Navajo Indians have been reported to be similar to those in other populations.[9] Thus, the high rates of NIDDM identified in the present study have significant clinical and economic implications for the health status and health care of Navajo Indians.

The large percentage of overweight individuals found in the population studied may contribute to the high prevalence of NIDDM. Because exercise programs for diabetes control have been successful in other Indian communities,[10] efforts at primary and secondary prevention by weight control and exercise are currently under way on the Navajo Reservation.

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Uncompensated Emergency Care in Hospital Markets in Los Angeles County

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Abstract: A survey of hospital emergency rooms in Los Angeles County was conducted in March 1987. Analysis of the distribution of uninsured emergency care patients revealed that private hospitals play a significant frontline role in terms of entry into the hospital system for patients who are unable to pay—almost one-half of such patients were treated in the emergency rooms of private hospitals. Hospitals serving markets in which a higher proportion of residents had incomes below the poverty level provided a greater share of uncompensated emergency room services. (*Am J Public Health* 1989; 79:514–516.)

Introduction

As public and private third party payors impose market-oriented reforms in hospital payment methods, concerns have heightened over the effects of these policies on the uninsured population. Attention has been directed to the access problems of the medically indigent under these new payment approaches and to the problems of disproportionate share hospitals serving the medically indigent. The imposition of competitive pricing methods limits the ability of hospitals to subsidize the costs of uninsured patients. Growth of competitive-based insurance programs has been particularly strong in California, where they now cover over 60 per cent of the population.[1] Nationally, this number is approaching 25 per cent.[2] We discuss the distribution of unsponsored hospital emergency room care in a large metropolitan area in California and explore whether hospital market and financial characteristics influence provision of care to uninsured patients.

Methods

A survey of hospital emergency rooms in Los Angeles County was conducted by the Los Angeles County Medical Association and the Hospital Council of Southern California for a one-month period, beginning in March 1987. The survey collected data on the volume and costs of emergency room care provided to patients who were both uninsured and unable to pay (hereinafter referred to as uninsured). The costs of treatment included hospital and physician costs associated with emergency room services as well as hospital inpatient costs for uninsured patients who were admitted from the emergency room.

Both public and private hospitals were surveyed. Hospitals owned by Kaiser-Permanente, a closed system health maintenance organization (HMO), were excluded from the analyses since they serve members enrolled in Kaiser health plans and treat few, if any, uninsured patients. All four County-owned acute care hospitals and 51 of the 81 private hospitals with emergency rooms responded.

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Hospital cost reports for fiscal year 1984 and discharge data for calendar year 1983 were provided by the California Office of Statewide Health Planning and Development. Emergency room charges for services rendered to uninsured patients as reported in the survey were reduced to costs using cost-to-charge ratios obtained from the hospital cost reports. We used patients' zip code of residence obtained from the discharge data to define each hospital's primary market area. The defined market area contained all zip codes that contributed at least 1 per cent of the hospital's discharges. If this rule did not result in a market area that contained at least 40 per cent of the hospital's discharges, additional zip codes were added (in descending order of patient volume) until a hospital's market area encompassed a minimum of 40 per cent of its discharges. We calculated the proportion of the population with income below the poverty level for each hospital's primary market area using 1980 census data.

Using multiple regression, we explored the role of several factors in explaining the wide variation among private hospitals in the volume and costs of care provided to uninsured patients. The independent variables included: ownership status, profit margin in a past year, hospital share of paying emergency room patients in the County, and the proportion of the population below the poverty level in a hospital's market area.

Results

Of all patients treated in the emergency room of responding hospitals, 19.8 per cent were classified as uninsured and unable to pay (see Table 1). A total of 36,304 such patients were treated in survey hospitals during the one month reporting period, incurring costs of $39.6 million. Although the four County hospitals provided treatment to the majority of uninsured patients, almost one-half of such patients were treated by the 51 private hospitals; the 15 private hospitals with trauma centers treated almost as many of these patients as the other 36 private hospitals.

Our estimates of the total number of uninsured emergency room patients and the total treatment costs incurred do not represent the total costs of treating such patients in the County since 35 per cent of hospitals with emergency rooms did not respond to the survey. Non-responding hospitals tended to be smaller than responding hospitals, having an average bed size of 198 compared to 312. All non-respondents consisted of private hospitals so the estimates of the relative burden of care borne by County-owned versus private facilities underestimate the contribution by private hospitals. However, we believe that this is not a significant problem since only one of the 16 trauma centers in the County were among the non-respondents. A telephone follow-up to a sample (eight hospitals) of the 30 non-respondents revealed that on average a low proportion of patients treated in the emergency rooms were uninsured and unable to pay.

On average two-thirds of the patients in County hospital emergency rooms are uninsured and unable to pay compared to 16 per cent in private hospital emergency rooms (see Table 2). Within individual County hospitals, the proportion of uninsured patients ranged from 48 per cent to 78 per cent, whereas within private hospitals, it ranged from less than 1 per cent to 42 per cent. Table 3 shows the distribution of uninsured patients among the hospitals.

TABLE 1—Distribution of Emergency Room Volume and Costs of Uninsured Patients Who Were Unable to Pay (UUTP), Los Angeles County, California, March 15–April 15, 1987

Hospital Type	# Responding Hospitals	# UUTP Patients	% of Total UUTP Patients*	% of Total Sponsored Patients	Total Cost	% of Total Cost
County	4	20,139	55.5	14.5	$31,783,114	80.3
Private	51	16,165	44.5	85.5	7,823,356	19.8
Trauma Center	15	7,474	20.6	36.6	3,924,016	9.9
Non-Trauma Center	36	8,691	23.9	48.9	3,899,340	9.8
All Hospitals	55	36,304	100.0	100.0	39,606,470	100.0

*This column refers to the proportion of all UUTP emergency room patients in the County that are treated by the hospitals in each group.

TABLE 2—Relative Burden of Uninsured and Unable to Pay Emergency Room Patients by Hospital Type, Los Angeles County, California, March 15–April 15, 1987

Hospital Type	Average Proportion of Uninsured Emergency Room Patients
	%
County	66.9
Private	16.1
Trauma Center	16.8
Non-Trauma Center	15.8
All Hospitals	19.8

Table 4 shows the estimated regression coefficients for each private hospital's share of the uninsured patients treated and for each hospital's share of the total cost of treating such patients.

The regression analyses show that hospitals with trauma centers provide a greater share of unsponsored care, both in terms of the number of uninsured patients and the total cost. Controlling for a hospital's share of paying patients, hospitals with trauma centers treat 1.6 per cent more of the uninsured patients and incur 2.1 per cent more of the costs of such patients on average than hospitals without trauma centers. Hospitals serving markets with higher poverty rates provide a greater share of services to this population. The estimated coefficient for a hospital's past profit margin is insignificant, suggesting that hospitals with higher profits in earlier periods do not use the surplus funds to provide a larger share of care to uninsured patients.

Discussion

Any solution to the problem of access to care for the medically indigent population needs to be broadly based and should address the complex issues surrounding access. Even though Los Angeles County has a large, publicly funded health care delivery system, private hospitals treat a substantial share (at least 45 per cent) of uninsured patients seeking emergency hospital care. Although the private hospitals transfer many of the patients to County-owned facilities, they play an important frontline role in providing entry into the health care system.*

County-owned hospitals treated 55 per cent of the uninsured emergency room patients but incurred 80 per cent of the total costs of treating such patients. The large share of costs borne by County facilities is likely due to several factors: 1) after providing initial treatment to uninsured patients, private hospitals often transfer these patients to County-owned facilities for further treatment; 2) County facilities may treat a more complex case mix of uninsured emergency room patients than do private hospitals; and/or 3) the average costs of County facilities are higher than in private hospitals.

Higher profit margins in a previous year do not lead to higher spending for care to uninsured patients, nor does a lower profit margin lead to lower spending. Unlike other parts of the United States, private not-for-profit hospitals in Los Angeles County provide only slightly more unsponsored care than do investor-owned hospitals.

*This may be particularly true in areas of Los Angeles County that are not near any of the County-owned facilities.

TABLE 3—Percentile Distribution of UUTP Patients within Hospitals, Los Angeles County, California, March 15–April 15, 1987

Percentage of Uninsured Emergency Room Patients	Number of Hospitals	
	Private	County
%	N	N
0–5	7	—
6–10	6	—
11–15	12	—
16–20	12	—
21–30	10	—
31–40	3	—
41–50	1	1
50	0	3
Average % Uninsured	16.1	66.9
Standard Deviation	0.09	0.14

TABLE 4—Determinants of Emergency Room Care to Uninsured and Unable to Pay Patients

Independent Variables	Dependent Variables: Share of Uninsured Patients	Dependent Variables: Share of Unsponsored Costs
Intercept	−0.014	−0.014
	(0.007)	(0.009)
Trauma Center	0.016	0.021
yes = 1, mean = .294)	(0.004)	(0.005)
Ownership	0.006	0.005
(nonprofit = 1, mean = .745)	(0.005)	(0.006)
Share of Emergency Room Paying Patients in County	0.272	0.263
(mean = .019)	(0.104)	(0.134)
Profit Margin	−0.013	−0.022
(mean = .064)	(0.031)	(0.040)
Poverty Percentage in Hospital's Market (mean = 13.0)	0.002	0.002
	(0.000)	(0.001)
R-Square	0.567	0.492

NOTES: The variable means are based only on the private hospitals in the sample since County-owned hospitals were excluded from the regression analysis.
Standard errors are in parentheses.

Changes in reimbursement policies in California (i.e., selective contracting by Medi-Cal and private insurers) are creating an increasingly competitive operating environment for hospitals. Hospitals located in areas with higher proportions of the indigent population will increasingly be at a competitive disadvantage. Recent attempts by hospitals in downtown Los Angeles to close their emergency rooms suggest that these pressures are already being felt.[3] Although the Medicare prospective payment system program provides extra payments to hospitals serving a disproportionate share of low income patients, this provision is scheduled to be discontinued in 1989. Emerging payment reforms need to recognize and incorporate provisions to ensure that access to medical care is not sacrificed in the effort to contain hospital costs.

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ADDENDUM

Background Information on Closure of Emergency Departments in Los Angeles as of September 1988

In May 1988, a private, not-for-profit hospital in downtown Los Angeles filed an application to "down-license" its emergency room from a full-service department to a standby facility that would no longer accept ambulances (it received approximately 800 ambulance visits per month). Soon afterward, three nearby hospitals filed similar applications, citing the potential increase in indigent patients if ambulances were re-routed to their facilities. Other hospitals have since threatened to follow suit.

As a short run solution, the County Department of Health Services provided $1.81 million in supplemental funding directly to eleven private hospitals that it believes are vital to the emergency services network. In return, the hospitals agreed to keep their emergency rooms open for the period of the contract. The County has agreed to provide an additional $1.25 million for a two-month period, but only on the condition that a statewide Tobacco Tax Initiative passes in the November election.

Of the eleven contracting hospitals, at least three had once been designated trauma centers but had withdrawn from the trauma network in 1987. Nine of the eleven are not-for-profit facilities, and the remaining two are investor-owned. Most are located in inner city areas. Absent a long run solution to the funding problem, the County could face a substantial reduction in the availability of hospital emergency department services.

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Television Viewing and Obesity in Adult Males

LARRY A. TUCKER, PHD, AND GLENN M. FRIEDMAN, MD

Abstract: We estimated the extent to which time spent watching television is associated with obesity and super-obesity among 6,138 employed adult males. After adjustment for age, smoking status, length of work week, measured physical fitness, and reported weekly hours of exercise, people who viewed TV more than three hours/day were twice as likely to be obese as those who viewed less than 1 hour/day. Those who viewed for 1 to 2 hours daily had a relative risk of 1.60 (1.21, 2.11). Physical fitness consistently confounded the associations between TV viewing and obesity/super-obesity, but the other control variables did not do so. (*Am J Public Health* 1989; 79:516–518.)

Introduction

Television viewing is the most pervasive pastime in the United States today. Following sleep and work, it is the nation's third most time-consuming activity.[1] The typical adult watches TV nearly four hours daily[1,2]; hence, it is not surprising that contemporary research indicates that human beliefs and practices are affected by television to a degree far exceeding earlier judgments.[1,3]

The role television plays in the development of health-related attitudes and behaviors is of growing interest to many.[4] Studies of the content of this powerful medium suggest that many health messages are conveyed regularly to viewers. Unfortunately, the information is sometimes unrealistic, distorted, and misleading,[7] particularly regarding food, nutrition, and obesity.[8–14]

Although many of the "micro-lessons" to which Americans are regularly exposed may promote misconceptions and produce unhealthy eating habits, television's primary offense may be one of omission rather than commission.[15] Research has shown repeatedly that the medium has profoundly altered American leisure.[1,16] When the TV is on, activity ceases and time for exercise is reduced significantly. The heart and other muscles of the body are not strengthened and calories are not expended in excess of resting metabolism during television viewing.

Recently, Tucker[17] examined the relation between television viewing and physical fitness. Results showed that as TV watching increased among 379 high school males, multiple measures of physical fitness decreased markedly and systematically. Similarly, Dietz and Gortmaker[18] showed that as TV viewing increased among several thousand children, obesity increased substantially.

The present study measured the extent of the association between TV viewing and obesity among adult males; an ancillary objective was to determine the extent to which age, cigarette smoking, physical fitness, time reported exercising, and hours worked per week mediate the relation between the television viewing and obesity.

Methods

Study subjects were 6,138 adult male employees of over 50 different companies that participated in the Health Examination Program offered by Health Advancement Services (HAS), Inc. Approximately 77 per cent of the subjects were married, 85 per cent were white, 51 per cent had some college education, and 32 per cent of the subjects were current smokers. The mean age

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was 39.5 (SD = 10.2) and the median and modal gross family income was $25,000–$30,000. Approximately 50 per cent of the subjects classified their jobs as professional or technical and 27 per cent as managerial/administrative.

All data were collected by registered nurses employed by HAS. Each subject was examined individually and privately for approximately 50 minutes after participating in a general orientation and completing an informed consent form. A written questionnaire was administered to assess demographic information, time spent watching television per day, number of cigarettes smoked per day, hours worked per week, and time spent exercising per week. Using a Harpenden skinfold caliper, subcutaneous fat was assessed at: *thigh* (anterior aspect midway between the hip and the knee-vertical fold), *chest* (half the distance between the anterior axillary line and the nipple-diagonal fold), *abdomen* (laterally at a distance of 2 centimeters from the umbilicus-vertical fold). The sum of the skinfold measurements along with age and gender were used to calculate the total body fat percentage of each subject. Physical fitness was assessed using a step test, the 3-minute Pulse Recovery Test.[19]

Obesity was treated categorically since body fat percentage poses little concern or risk until elevated levels signifying excessive fat are attained. Specifically, obesity was defined as 21 per cent to 30 per cent body fat[20]; males with 31 per cent or more body fat were considered super-obese. Television viewing and the control variables were categorized as shown in Table 1.

The relation of obesity and super-obesity to television viewing was measured by the odds ratio[21] with infrequent television viewers as the reference group. Mantel-Haenszel summary risk estimates were used to control for potential confounders. The relation of the control variables to television viewing and obesity was measured by the Mantel-Haenszel chi-square statistic.[22–24]

Results

As shown in Table 1, most subjects reported one to two hours of television viewing per day: more television viewing was reported by younger subjects, by the less fit, by smokers, by non-exercisers, and by subjects working fewer hours. Obesity or super-obesity, measured in nearly one-third of the subjects, varied markedly according to age, measured fitness, and reported exercise but not with smoking status or length of work week.

Table 2 shows the estimated relative risk of obesity by television viewing time without adjustment, with adjustment for all of the control variables identified in Table 1, and with adjustment only for measured fitness, the only actual confounder. After adjustment, subjects who viewed three hours of television or more each day showed twice the risk of obesity and even greater risk of super-obesity.

Discussion

The present findings indicate that duration of daily television viewing is strongly associated with obesity and super-obesity in adult males, as in children.[18] The associations could mean that: obesity causes increased television viewing; that obesity and television viewing are each caused by other common factors; or that increased television watching causes obesity.

By the first interpretation, men who are fat may be attracted to watching television as a primary source of recreation or entertainment, whereas non-obese men may prefer other pastimes, since most leisure activities require more physical exertion than television watching.

If the second interpretation is correct, the data suggest that the common correlate confounding the association is not age or length of work week. Smoking may confound among adolescents,[25] but it did not confound in this adult population.

The confounding effects of time reported exercising and physical fitness on the TV viewing/obesity relationship have not been studied previously; in these data, both were inversely related to television viewing and to obesity. Adjustment for time reported exercising (a subjective measure) reduced the relative risk estimates by only 4 per cent on the average, while adjustment for measured physical fitness reduced the estimates by 13 per cent for obesity and 28 per cent for super-

TABLE 1—Television Viewing Hours and Obesity, According to Control Variables

		Television Hours/Day (%)				Obesity (%)		
	N	<1	1–2	3–4	>4	Non-Obese	Obese	Super Obese
Total Group	6,138	9.5	61.0	25.3	4.2	69.0	28.3	2.7
Age (years)								
19–29	970	10.1	54.3	29.7	5.9	83.6	15.1	1.3
30–39	2,243	8.7	60.7	26.2	4.4	72.7	25.2	2.1
40–49	2,028	10.0	65.5	21.9	2.6	66.4	31.5	2.1
50–59	697	9.0	64.8	23.0	3.2	51.4	43.2	5.4
60+	200	6.0	64.8	24.5	4.7	47.6	43.3	9.1
Physical Fitness								
good	457	15.4	61.2	22.1	1.3	90.3	9.0	.7
average	2,927	9.8	61.6	23.7	4.9	75.1	23.6	1.3
poor	1,805	7.5	59.1	28.6	4.7	55.8	40.2	4.0
couldn't complete	142	6.0	53.0	31.6	9.4	41.9	50.8	7.3
did not participate	807	5.6	57.1	31.9	5.4	47.5	44.6	7.9
Smoking Status								
smoker	2,006	6.2	57.1	31.1	5.6	69.1	28.6	2.3
non-smoker	4,132	10.7	64.2	22.0	3.1	68.5	28.6	2.9
Exercise Status								
mild/non-exerciser	3,356	8.5	58.8	27.8	4.9	65.2	31.6	3.2
regular exerciser	2,782	10.3	65.3	21.5	2.9	75.2	23.0	1.8
Work Week								
40 hr or less/wk	2,041	7.9	56.7	29.4	6.0	70.5	26.8	2.7
41–50 hr/wk	3,157	9.0	64.0	24.1	2.9	69.3	28.2	2.5
51 hr+/wk	940	13.2	66.9	17.5	2.3	64.7	32.3	3.0

TABLE 2—Relative Risk of Obesity and Super-obesity by Television Viewing Group

Daily TV Viewing	Variable Controlled	Obese				Super-obese			
		N	%	RR_{mh}	95% CI	N	%	RR_{mh}	95% CI
<1 hour									
(*n* = 584)	none	110	18.8	1.0*	—	12	2.1	1.0*	—
1–2 hours									
(*n* = 3,744)	none	1051	28.1	1.69	1.36, 2.10	82	2.2	1.21	0.66, 2.23
	age, fitness, smoking, exercise & work			1.60	1.21, 2.11			1.08	0.51, 2.28
	physical fitness only			1.62	1.26, 2.08			1.06	0.55, 2.03
3–4 hours									
(*n* = 1,551)	none	488	31.5	2.05	1.62, 2.58	61	3.9	2.34	1.27, 4.32
	age, fitness, smoking, exercise & work			2.05	1.48, 2.84			2.33	1.18, 4.63
	physical fitness only			1.76	1.35, 2.30			1.73	0.90, 3.33
>4 hours									
(*n* = 259)	none	88	34.0	2.34	1.68, 3.25	13	5.0	3.17	1.47, 6.83
	age, fitness, smoking, exercise & work			1.90	1.06, 3.38			*	—
	physical fitness only			1.87	1.29, 2.69			1.69	0.67, 4.29

NOTE. RR_{mh} = Mantel-Haenszel summary estimate of relative risk.
*Unable to calculate because of insufficient subjects in the frequent viewers super-obese subgroup and the many strata of the control variables.

obesity. Nonetheless, television time and obesity were strongly related even after holding fitness constant.

If the third interpretation is correct, a possible mediator, unmeasured in the present study but equal in theoretical importance to time spent exercising and physical fitness, is caloric intake. Because television viewers are bombarded by thousands of messages for nonnutritious foods via advertisements and prime-time programs,[9,10,14] and because television viewing correlates with snacking and consumption of foods advertised on television,[7,26,27] it is possible frequent viewers consume significantly more calories than their counterparts, and hence are more obese.

Future research on the association of television viewing and obesity will need to consider this factor. In addition, because subjects in this study had relatively high SES (socioeconomic status), generalization to poorer groups will require additional research.

With the growth of cable television, home video recording, and videogames, television viewing is likely to increase in the coming years. The findings of this study and other recent research show that the impact of television on the lifestyles and health of Americans cannot be ignored.

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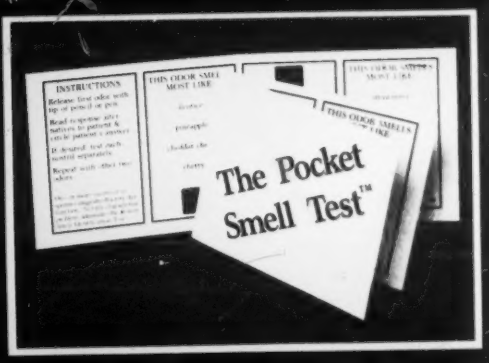

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Trends in Dental Fluorosis and Dental Caries Prevalences in Newburgh and Kingston, NY

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Abstract: A study was undertaken in New York State to determine the changes in dental fluorosis prevalence from 1955 to 1986 in fluoridated Newburgh and non-fluoridated Kingston children. The frequency and severity of dental fluorosis among 884 7–14-year-old children were measured by two dentists utilizing Dean's Index. Data regarding residential and fluoride history were obtained from the parents of participants. Among the Newburgh residents, the prevalence of dental fluorosis (very mild to moderate) varied from a low of 5 per cent for the 9–10-year-old group to a high of 9.4 per cent for 11–12-year-olds. Except for the 13–14-year-old group, children in non-fluoridated Kingston had the lowest dental fluorosis prevalence rates. A comparison of Dean's Community Fluorosis Indices to the 1955 baseline data obtained from studies conducted after 10 years of fluoridation in Newburgh revealed no changes of consequence among Newburgh residents. However, the changes are apparent for Kingston residents, indicating the availability of fluorides in non-fluoridated areas. The increased risk for dental fluorosis for Kingston residents appears to be from the use of fluoride tablets. An analysis of dental caries data revealed that caries prevalence declined substantially in both fluoridated and non-fluoridated areas. (*Am J Public Health* 1989; 79:565–569.)

Introduction

Oral health surveys conducted in the United States and other developed countries have demonstrated that the prevalence of dental caries in children has declined substantially in recent years.[1–8] The increased availability of fluoride in many forms appears to be the single most important factor. The use of antibiotics, improved levels of restorative care, and patterns of sugar consumption and snacking have also been thought to contribute to the secular changes.[1,9–11]

To explain changes in dental caries prevalence in non-fluoridated areas, many investigators have pointed out the effect of systemic fluorides in professionally prescribed supplements and/or the use of fluoridated dentifrices by young children.[1,12–15] The increased availability of fluorides in food has also been suggested.[13,16–21] Singer and his co-workers have indicated that the average daily fluoride intake for infants, two-year-olds, and young male adults is within the accepted range of 0.05 to 0.07 milligram per kilogram body weight and has not significantly changed over the past 30 years.[22–25] However, studies of fluoride levels of baby formulas and cereals have shown a significant increase in the fluoride content when fluoridated water was used for processing these foods.[12,19,27–29]

Additional sources of fluorides are available now that were almost nonexistent when water fluoridation standards were determined. These include fluoride rinses and fluoride dentifrices. Barnhart, *et al.*[30] reported that 2- to 4-year-old children ingested an average of 0.3 gram (gm) of dentifrice or 0.3 milligram (mg) of fluoride per brushing. Fluoride supplements are commonly prescribed for children living in non-fluoridated areas.

The availability of fluoride from multiple sources has raised questions about the potential for increases in the incidence of dental fluorosis. Aasenden and Peebles[31] found that 67 per cent of the children who ingested 0.5 mg fluoride from birth to three years of age and 1.0 mg daily thereafter had very mild to mild enamel fluorosis and 14 per cent had moderate fluorosis. In a later report, they suggested that fluorosis might become less noticeable with age.[32] Hennon and co-workers[33] reported no unacceptable levels of fluorosis among children who resided in communities with 0.6 to 0.8 ppm fluoride in the water and had taken 0.5 mg fluoride supplements from birth to age 3 years. Recently, Soparkar, *et al.*,[34] reported an unexpectedly high proportion of children with enamel fluorosis and the source of fluoride was determined to be prescribed fluoride supplements. Oldak and Leverett[35] observed an increased occurrence of dental fluorosis in children residing in non-fluoridated communities. Twenty-two per cent of children ages 6–8 residing in a non-fluoridated community had fluorosis in permanent teeth. This higher than expected level of dental fluorosis was thought to be due to the prior use of dietary fluoride supplements.

Recently, Driscoll, *et al.*,[36,37] and Segreto, *et al.*,[38] compared data on the prevalence of fluorosis with historic data and found no important changes in the prevalence and severity of fluorosis. Although mean fluorosis indices were within the normal range in Rochester, New York, a higher than expected prevalence of very mild categories of fluorosis was reported by Leverett.[39]

Most epidemiological studies of dental fluorosis have been conducted in the Midwestern States, but baseline data on the frequency and severity of dental fluorosis in New York State are available from the 1955 Newburgh-Kingston Study.[40] Therefore, an opportunity existed to determine if the frequency and severity of dental fluorosis had changed over the last 30 years in those two cities.

Similarly, changes in dental caries prevalence could also be determined. Studies conducted between 1944 and 1955 in Newburgh indicated that dental caries prevalence declined considerably after fluoridation. In the non-fluoridated Kingston, dental caries prevalence increased during the same period. Another study of 13- and 14-year-old children in Newburgh and Kingston revealed that this trend continued into the early 1960s.[40–42]

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Methods

Between 1955 and 1986, several demographic and socioeconomic changes have occurred in both Newburgh and Kingston. Compared to Kingston, Newburgh has: a lower median income, a slightly larger proportion of low income level households, a lower median school years completed, a lower percentage of White population, a slightly higher percentage of unemployed, and a different occupational distribution.[43] The water supply records in Newburgh indicate that the level of fluoride in the water was maintained at the recommended 1 ppm established by the US Public Health Service in 1945, except for a three-year period from 1978 to 1981. This reduction in fluoride would affect the teeth of 7- to 14-year-olds differentially, depending on the stages of development of the teeth during this time. The city of Kingston has a fluoride content of 0.3 ppm in water.[44]

To obtain estimates of prevalences of dental fluorosis and dental caries in the cities of Newburgh and Kingston, children attending grades one through eight in public schools were selected to participate on the basis of a stratified random sampling of clusters. A total of 1,446 children from five elementary and two junior high schools in Newburgh and 694 children from four elementary and two junior high schools in Kingston agreed to participate in clinical examinations. The response rates were 63 per cent for Newburgh and 50 per cent for Kingston. The overall response rates were 72 per cent and 38 per cent for elementary and junior high schools, respectively.

Dental fluorosis was determined for each child according to the established classification by Dean, in which the child is classified on the basis of the two teeth in the mouth showing the most advanced signs of fluorosis.[45] In order to differentiate fluorosis from other enamel defects, the criteria developed by Russell was used.[46] Dental caries was assessed using World Health Organization (WHO) guidelines.[47] A pretested questionnaire was used to obtain residential and fluoride history from the parents of participants.

Two dentists, who were not involved in New York State fluoridation or fluoride supplement programs, independently examined approximately equal proportions of participants in Newburgh and Kingston. Examination equipment included a portable chair and a fiber optic light source. The examiners were trained by an outside consultant in applying Dean's classification and were standardized four times during the study. Further, as an attempt to incorporate blindness, children who were not continuous residents of Newburgh and Kingston were also examined. The residential and fluoride history data were not made available to the examiners.

The analysis of data is limited to groups of children between ages 7 and 14. Children under age six and over age 15 were excluded because of inadequacies in the sample size. Children with orthodontic bands or only deciduous teeth were also excluded. Estimates of community fluorosis indices and respective standard errors were calculated for comparison purposes using the methods appropriate for stratified cluster sampling.[48] Trends in dental caries prevalences were determined by comparing age-race adjusted decayed, missing, filled, teeth (DMFT) indices. The relationship between fluorosis and dental caries was determined by analysis covariance technique. The adjusted mean scores were derived by treating age as a covariate utilizing the SAS-GLM computer package.[49]

Because a large number of children attending the schools in Newburgh City were from the non-fluoridated town of New Windsor and the recently fluoridated town of Newburgh, the children were separated into distinct groups. As the amount of exposure to fluorides is of particular importance, data analysis is limited to continuous residents of Newburgh and Kingston. For the purpose of this report, 993 children who were not lifetime residents of Newburgh or Kingston cities were excluded from the analysis. Because the prevalence and severity of fluorosis did not differ significantly by ethnic group, sex, or examiners, only summary data are presented. As age-specific fluorosis rates tended to be unstable, data are presented by two-year age intervals.

Results

The characteristics of the study population by relevant variables are presented in Table 1. There are smaller proportions of White and a generally lower set of socioeconomic indicators in the Newburgh sample and somewhat fewer lifetime residents (14 per cent vs 16 per cent).

The frequency and severity of dental fluorosis for the children examined is presented for the two community groups in Table 2. Among the residents of the fluoridated Newburgh City, as well as residents of Kingston, the majority of the detectable dental fluorosis lesions observed were in the very mild category.

Among the Newburgh residents, the prevalence of dental fluorosis (very mild to moderate categories) varied from a low of 5 per cent for the 9–10-year-old group to a high of 9.4 per cent for 11–12-year-olds. Except for the 13–14-year-old group, children in the non-fluoridated Kingston had the lowest dental fluorosis prevalence rates. The relation between age and dental fluorosis prevalence was inconsistent.

When the distribution of the lesions on the worst affected tooth of each child was analyzed, the clinical manifestation of the lesions observed in Newburgh and Kingston was similar (Table 3).

A comparison of Dean's community fluorosis indices among the populations reveal that there are no differences in 1986 between Newburgh and Kingston residents (Table 4). Table 4 also compares the Dean's community fluorosis indices to the baseline data obtained from 1955 studies conducted after 10 years of fluoridation in Newburgh and Kingston. Although changes in the prevalence of dental

TABLE 1—Comparison of Demographic and Socioeconomic Characteristics in Newburgh and Kingston, NY

Characteristics	Newburgh	Kingston
Number of 7–14 year-old Children	3,209	2,611
Per Cent White	47.5	84.8
Median Family Income	$14,230	$16,573
Per Cent Unskilled (Employed)	24.1	22.0
Per Cent with 4 or More Years of College	7.7	13.0
Low Income Households in Per Cent (Yearly Income Less than $4,000)	25.1	18.2
Employed Persons 16 years and Over in Per Cent	48.1	51.5
Lifetime Residents	459	425
Per Cent Male	47.7	50.3
Per Cent White*	33.2	79.7

Source: US Bureau of the Census: 1980 Census of Populations, General Social and Economic Characteristics, New York, PC80-1-C-34NY.

*NOTE: The Black, Hispanic and Asian children were categorized as Other than White population; 78.5% and 79% of the Other than White population were Blacks in Newburgh and Kingston, respectively.

TABLE 2—Percentage Distribution of Dental Fluorosis in Newburgh and Kingston Residents by Age Group and Dean's Classification

Age	Number	Dean's Fluorosis Classification				
		Questionable	Very Mild	Mild	Moderate	Prevalence
Newburgh						
7– 8	149	11.4	4.6	2.6	2.0	9.2
9–10	135	12.5	2.2	2.9	0.0	5.1
11–12	104	12.5	7.6	0.9	0.9	9.4
13–14	71	25.3	5.6	1.4	0.0	7.0
Kingston						
7– 8	120	11.6	4.1	0.8	0.0	4.9
9–10	110	11.8	1.8	1.8	0.9	4.5
11–12	101	13.8	3.9	3.9	0.9	8.7
13–14	94	7.4	8.5	2.1	1.0	11.6

NOTE: Dental fluorosis prevalence is based on very mild to moderate categories of fluorosis. Children classified as normal or questionable are viewed as not having dental fluorosis.

TABLE 3—Distribution of Fluorosis in 7 to 14 Year-Old Children in Newburgh and Kingston According to Clinical Manifestation

Fluorosis	Newburgh %	Kingston %
Generalized Lesions	83	76
Bilateral Symmetry	96	97
Affected Areas (Cuspal Tips Only)	88	81
Degree of Tooth Affected (Less than 25% of the Tooth)	85	82

TABLE 4—Comparison of Community Fluorosis Indices Along with Standard Errors among Groups by Age and Year of Examination

Age (years)	Newburgh		Kingston	
	1955	1986	1955	1986
7– 8	0.18 (.04)	0.21 (.05)	0.00	0.16 (.04)
9–10	0.14 (.02)	0.14 (.02)	0.00	0.13 (.03)
11–12	0.11 (.03)	0.20 (.07)	0.00	0.23 (.05)
13–14	0.00*	0.20 (.05)	0.00	0.19 (.05)

*Newburgh children of this age group did not receive fluoridated water during the formation of teeth.

NOTE: The 1955 Community Fluorosis Indices for Newburgh and Kingston residents were calculated from the frequency and severity of Dental Fluorosis.[40]

fluorosis are not apparent among Newburgh residents, changes for Kingston residents were detected.

A history of the use of fluoride tablets/drops during the first eight years of life from the parental questionnaire may indicate the exposure to a known risk factor. Table 5 shows that approximately 31 per cent of the children in Kingston reported a history of regular use of fluoride tablets/drops compared to 5 per cent for the residents of Newburgh. The odds ratio calculated for the two groups indicated a greater odds of developing dental fluorosis when tablets were taken every day. Over 95 per cent of the children in both groups reportedly use fluoride dentifrices. The effect of other sources of fluoride, such as topical rinses and professional fluoride applications was not examined.

A comparison of the data gathered from studies conducted in Newburgh and Kingston reveals that caries has declined both in fluoridated and non-fluoridated areas (Figure 1). Figure 2 demonstrates the caries trends in 13- and

TABLE 5—History of Fluoride Tablets/Drops Use and the Estimated Relative Risk with Everyday Use of Tablets/Drops during the First Eight Years of Life

	N	Fluoride Tablet/Drops History			Odds Ratio*
		About Every Day	About 3 Days/Week	Occasionally	
Newburgh	459	24 (5%)	1 (0%)	47 (10%)	1.7
Kingston	425	133 (31%)	16 (4%)	46 (11%)	3.8

*Indicates the ratio of the prevalence of dental fluorosis among those who reported everyday tablet/drops use compared to those who did not report tablets/drops use.

Odds Ratio (Newburgh) = 3/24 ÷ 28/387 = 1.7 (95% CI = −0.5, +3.9)

Odds Ratio (Kingston) = 20/133 ÷ 9/230 = 3.8 (95% CI = +.07, +6.9)

14-year-old children.

A comparison of dental fluorosis and dental caries prevalence suggests an inverse relationship. Children with very mild to moderate dental fluorosis consistently had lower caries experience (Table 6). The covariate adjusted mean DMFS indices showed that children with very mild to moderate fluorosis had approximately 41 to 54 per cent fewer DMFS, respectively.

Discussion

Although demographic and socioeconomic changes over a period of four decades have resulted in population differences between Newburgh and Kingston, an analysis of trends in fluorosis prevalence should be still valid. Segreto, *et al*, have reported that the effect of gender, ethnic group, or family income in dental fluorosis is small compared with the effect of optimal fluoride level in drinking water.[38] The changes in dental caries prevalence in Kingston must be viewed in light of the effect of changes in socioeconomic characteristics and the impact of the availability of other forms of fluoride in non-fluoridated areas.

The results indicate that the extent and severity of dental fluorosis among residents in an optimally fluoridated community in a Northeastern State is within the normal range, as specified by Dean. Dean has stated that the mildest forms of dental fluorosis would occur in about 10 per cent of a population that used water containing about 1 ppm of fluoride continuously from birth.[50] Later, he suggested a method to

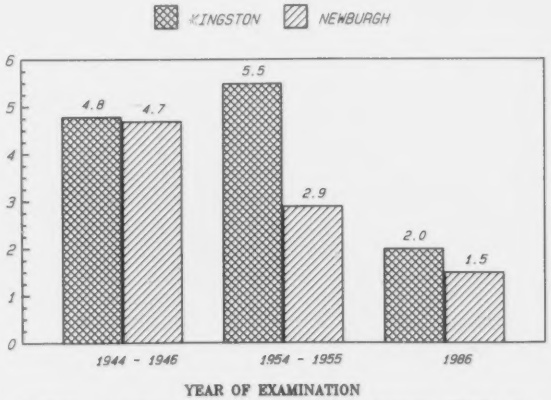

FIGURE 1—Age-Race-Adjusted Dental Caries Prevalence (DMFT) for 7–14-Year-Old Children in Newburgh and Kingston, NY, 1944–86

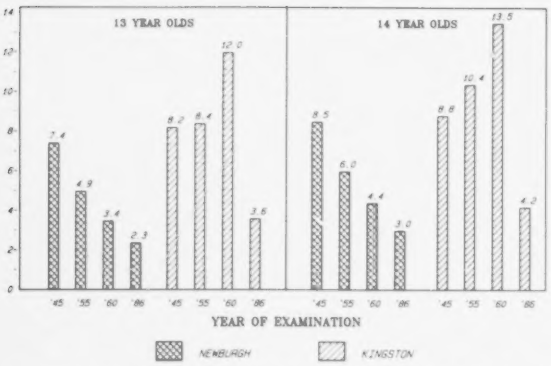

FIGURE 2—Dental Caries Prevalence for 13- and 14-Year-Olds in Newburgh and Kingston, NY, 1944–86

ascertain the public health significance of dental fluorosis utilizing the Community Fluorosis Index (CFI). According to Dean, CFI scores below 0.4 are considered as having no public health significance.[51] Because 99 per cent of the population studied had less than moderate categories of dental fluorosis and the differences over time were negligible, these data suggest that no changes have occurred in the past three decades. This observation is consistent with the recent studies conducted by Segreto, *et al*, in Texas,[39] and by Driscoll, *et al*,[36,37] in Illinois.

The lower community fluorosis indices for all age groups for Newburgh residents compared to those reported elsewhere[36–39] for residents of other fluoridated communities could be due to three years of interruption in fluoride availability from water. However, an analysis of the relation between age and fluorosis prevalence did not reveal any consistent effect of this interruption. The 13–14-year-old group that received an uninterrupted supply of fluoride from birth for seven years did not have any more fluorosis than the 7–8-year-old group that received the least amount of fluoride exposure from water. Also, the fact that 16 per cent of Newburgh resident children also reportedly received fluoride tablets may indicate that some of these children received fluoride during the period of interruption in water fluoridation. Further, compared to other indices that consider the tooth as a unit, the use of Dean's Community Fluorosis Index based on the child as a unit minimizes the underestimation resulting from interruption in fluoride availability, because any excessive exposure during the development of teeth will affect at least some teeth. This manifestation is considered as fluorosis, no matter whether the child received fluoridation during the development of some teeth or all the teeth.

Another possible explanation for the low community fluorosis indices may be due to differences in the application of the diagnostic criteria. This may have been minimized by not providing the fluoride history to examiners and by the high level of agreement among the examiners. A comparison to an earlier study in Newburgh and Kingston conducted by Russell in 1955 should take into consideration the differences that could have occurred in the application of the diagnostic criteria. The community fluorosis indices reported in 1955 for Newburgh residents were also lower compared to that of other studies.[45] However, at that time, only 7–8-year-old children had received fluoride in drinking water continuously from birth.

TABLE 6—Dental Caries Prevalence According to Dean's Fluorosis Classification

Dean's Fluorosis Classification	N	Covariate (Age) Adjusted Mean DMFS	Standard Error
Normal	704	2.2	0.11
Questionable	113	2.1	0.27
Very Mild	41	1.3	0.47
Mild	19	1.1	0.67
Moderate	7	1.0	1.11

The change in the occurrence of very mild to mild categories of dental fluorosis from the original study is noticeable in the non-fluoridated city of Kingston. This increased occurrence of fluorosis in non-fluoridated areas has been reported by other investigators.[4,34,35]

In Kingston, the increased occurrence of fluorosis may be attributed to a large extent to the use of fluoride supplements. If the use of infant formula was similar in the studied populations and the cases and controls are representative of the population, then an important factor associated with dental fluorosis is the use of fluoride tablets/drops. The fact that this group received fluoride supplements at a higher than the currently recommended dosage may have contributed to an increased risk for developing fluorosis.[52]

The higher dental fluorosis rate reported for Kingston when compared to the studies conducted in other non-fluoridated areas is consistent with the variation in the fluoride supplement use in different regions of the country.[34–39] The findings from the 1983 National Health Interview Survey indicated that the use of fluoride supplement was more frequent among White families with high incomes in the Northeastern regions.[53]

Dental caries declined substantially both in Kingston and Newburgh, a trend that is consistent with the recent findings from other studies.[1–9,13] An observation of particular interest is that between the 1940s and 1960s, dental caries increased in the non-fluoridated Kingston in contrast with the fluoridated Newburgh. As a result of this phenomenon, the caries decline between the 1960s and the 1980s in Kingston appears to be greater in Kingston compared to Newburgh. In spite of this decline in caries in the non-fluoridated area, lower caries prevalence continues to exist in the fluoridated area. However, factors such as differences in the service utilization patterns as evidenced by the comparison of the proportion of filled component of all caries surfaces, interruption in the fluoride availability for Newburgh children, and an increased availability of fluorides for Kingston children complicate the measurement of impact of fluoridation in this study.

An analysis of dental caries prevalence, according to Dean's fluorosis classification, revealed that children with very mild to moderate to dental fluorosis had 41 to 54 per cent lower caries prevalence. Driscoll, *et al*, observed similar differences among children residing in areas where the fluoride is above optimal levels.[37] The additional caries inhibitory effect for children with very mild to moderate dental fluorosis observed in the original studies appears to still exist.[51] Several researchers have suggested that the logical implication of an inverse relationship between dental caries and dental fluorosis in the development of public policy is that a slightly increased risk for dental fluorosis is preferable to a higher risk for dental caries,[57,58] since the very mild to mild categories of fluorosis are neither an adverse health effect nor esthetically objectionable, whereas dental caries results in pain, suffering, loss of esthetics, and higher costs.

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Survey of Cause-of-Death Query Criteria Used by State Vital Statistics Programs in the US and the Efficacy of the Criteria Used by the Oregon Vital Statistics Program

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Abstract: A survey of the 52 vital statistics registration areas in the United States revealed that at least 23 did not fulfill the minimum cause-of-death query guidelines recommended by the National Center for Health Statistics. The Oregon Center for Health Statistics is one of only a few that query certifying physicians at a comprehensive level. During August 1986–July 1987, a total of 2,453 of 23,238 death certificates were returned to the certifiers for additional information, not including those returned in a tobacco use study. More than one-half (56.1 per cent) resulted in new and more specific underlying cause-of-death data. Only 5.2 per cent of the queries were unanswered.

One probable result of Oregon's program is that the state has the highest percentage of liver cirrhosis and disease deaths attributed to alcohol abuse in the United States. Nationally, 41.7 per cent of all liver disease and cirrhosis deaths in 1984 were listed as due to alcohol compared to 82.4 per cent in Oregon. The state's total liver cirrhosis and disease death rate (12.0 per 100,000 population) is only marginally higher than the United States rate (11.6). The query program also serves to locate maternal deaths that would otherwise not be reported, as well as to provide more accurate cause-of-death statistics in general. (*Am J Public Health* 1989; 79:570–574.)

Introduction

Health statistics play an important role in revealing the existence and extent of health problems, both locally and nationally. Perhaps the most fundamental measure of the health status of any population is the life-death dichotomy. Mortality statistics are widely available in the United States for various populations but, since the vital statistics system is state-based, are the data collected in a comparable manner?

The underlying cause of death (the disease or injury that started the sequence of events leading directly to death) is a difficult concept for some physicians. These physicians often report the cause of death solely as a mode of death such as cardiac arrest or respiratory failure. The querying of certifying physicians by state vital statistics offices has long been recognized as an important method of improving the accuracy of statistics generated from death certificates; it also serves to educate physicians on how to properly record the cause of death.

In the United States, death certificates are collected by the 50 states plus two other vital statistics registration areas* (New York City and the District of Columbia). The information on cause of death reported by the certifying physician is coded by either the state nosologists, or by nosologists at the National Center for Health Statistics (NCHS) from copies of the death certificates. The NCHS is entirely dependent on the states to query physicians regarding certificates that were not completed satisfactorily. In June 1985, the NCHS published recommended criteria for cause-of-death querying.[1]

NCHS has described six priority levels for querying, with Level 1 consisting of the minimal amount of querying necessary "to promote the basic integrity of state and national data," and Level 6 being the most comprehensive level.[1] The six levels were developed recognizing that the availability of staff and resources vary from state to state but that Level 1 items should always be queried. Level 1 causes to be queried include entries such as "heart disease," "adult respiratory distress syndrome," and "senility." Each higher priority level provides additional detail. For example, the specific site of a malignant neoplasm of the pancreas is queried at Level 5.

This report summarizes the results of the survey of cause-of-death querying practices in state vital statistics programs. It also describes the procedures used by the Oregon Center for Health Statistics (OCHS) to assure that mortality statistics are as accurate as possible and the effect of the procedures on cause-of-death rates.

Methods

Oregon has had a cause-of-death query program for many years. Prior to 1983, nosologists manually selected certificates requiring further information. Beginning in 1983, query criteria were formalized, with a list of most certificates requiring queries generated by computer. Three months after the NCHS published recommended criteria for cause-of-death querying in June 1985,[1] the OCHS implemented a revised automated query generation system based upon NCHS guidelines. A small number cannot be selected automatically and require individual selection by nosologists. A separate list is generated, for example, when causes suggestive of AIDS (acquired immunodeficiency syndrome) (Appendix A) are listed but AIDS itself is not mentioned on the death certificate.

The OCHS chose to select certificates for query at priority Level 4, the most thorough level recommended for general querying. Levels 5 and 6 are "frequently required for specified special studies."[1] Additionally, the OCHS queries physicians on a few Level 5 criteria and for some causes not listed by the NCHS (see list in Appendix B). The OCHS no longer queries for some format problems, congestive heart failure, or mental conditions as recommended by the NCHS because, in our experience, these queries drew only useless or irate responses from physicians.

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*Ensuing references to "states" include New York City and the District of Columbia.

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The OCHS uses 27 basic formats for query letters. Each letter measures 21.6 × 12.7 cm and is photocopied adjacent to the death certificate in question, resulting in a legal-sized single sheet of paper containing a copy of the certificate and query form. This is then customized, if necessary, and mailed to the certifier along with a flyer explaining the query program and highlighting common problems requiring queries. The letters used by the OCHS are more specific to the cause in question than the sample letters developed by the NCHS.

Each query is subsequently categorized in one of five groups: 1) No response, 2) Responded with inadequate information, 3) Responded with adequate information but no changes, 4) Responded with adequate information resulting in a change to the underlying cause-of-death, 5) Responded with adequate information not resulting in a change to the underlying cause but resulting in additional cause-of-death information (such as substance abuse which is tracked with supplemental codes).

To compare Oregon's query procedures with those of other states, a questionnaire was mailed to each state registrar on October 20, 1986. Three questions were asked: 1) Do you use Part 20 of the NCHS Instruction Manual? 2) If so, what priority level do you use? 3) If not, what other criteria do you use? After a follow-up letter, all state registrars except those from Maryland, Massachusetts, Virginia, and New York City had returned their questionnaires.

Results

The Oregon Program

We report the outcome of 2,453 cause-of-death queries made on 23,829 certificates during August 1986–July 1987. These do not include tobacco use queries which are discussed later.

The response rate by physicians to queries is very good, 94.8 per cent. This includes 0.4 per cent who responded, but with inadequate information, and 34.8 per cent who responded with adequate information that did not result in a change to the underlying cause-of-death code. However, some of the queries in this category yielded new supplemental codes and additional external cause-of-death information. Beginning in January 1987, a new code was created to track queries that resulted in changes to the certificate but that did not affect the underlying cause-of-death. Six per cent of the 1987 certificates were so categorized (3.5 per cent during the entire study period). More than one-half of the queried items (56.1 per cent) resulted in more specific underlying cause-of-death data. This includes 1.0 per cent of the queried certificates that were recoded as a result of the nosologist reading selected medical examiner files.

Response by Type of Query Form

The most commonly used form accounts for 40.4 per cent of the total. It asks for the underlying cause of the stated condition; 49.0 per cent of these queries generate responses that result in a change of the underlying cause-of-death. The three most frequently queried causes are: *cardiac arrest* (with 55.7 per cent ultimately reassigned to the underlying cause), *septicemia* (49.4 per cent), and *renal failure* (54.4 per cent). Only two other forms are used more than 5.0 per cent of the time; one asks if the *primary site of a carcinoma* with an unspecified site was known, while the other is a letter customized for individual certificates with *conflicting or missing information*. Both result in more specific and/or accurate causes of death in about 55 per cent of the cases.

Eight of the forms result in new cause-of-death codes at least 80.0 per cent of the time. Queries regarding motor vehicle and watercraft accidents generated the largest percentage of new underlying cause-of-death codes. The form that generates the lowest percentage of changes in the underlying cause-of-death asks, "Was this woman pregnant within 42 days prior to her death?" Nonetheless, this is how the OCHS locates many maternal deaths, because not all physicians note the fact on death certificates. Female deaths meeting the criteria listed in Appendix B are queried.

In May 1987, the OCHS began querying certifying physicians to ascertain if tobacco use contributed to the death of the decedent when one of the following conditions was listed as the cause-of-death: lung cancer, bronchitis, emphysema, and chronic obstructive pulmonary diseases (COPD). Tobacco use was chosen for query because the death rates for these causes were 10.9, 33.3, 55.9, and 22.2 per cent higher, respectively, for Oregon than for the nation during 1985; Oregon has had elevated death rates for these causes for several decades.

Six hundred ninety-one deaths from these causes were recorded from May to August and 427 queried because the physicians failed to note the role of tobacco, if any. Only 3.7 per cent of the physicians failed to respond. Tobacco use was reported to have contributed to, or caused, the following deaths: lung cancer, 76.9 per cent; bronchitis, 70.0 per cent; emphysema, 81.8 per cent; COPD, 76.6 per cent. The OCHS records this information with a supplemental code because International Classification of Diseases (ICD)[2] and NCHS rules[3,4] do not permit the underlying cause-of-death to be coded to tobacco use when the tobacco has been reported as the cause of the above conditions. Prior to querying, tobacco use was mentioned on only 8.5 per cent of the certificates when the underlying cause was one of the four causes listed above. A question regarding tobacco use was added to the Oregon death certificate in January 1989 as a consequence of the data generated by the query program. It applies to all deaths, not just lung cancer and chronic obstructive pulmonary diseases.

Changes in Death Rates

Table 1 illustrates the effect of an ongoing query program on cause-specific death rates. The principle result of querying is the reassignment of codes from less specific categories to more specific ones and from conditions that do not usually arise spontaneously to their underlying cause. Six of the 20 leading causes of death in Oregon had death rates which were more than 10 per cent higher or lower after querying. Although the overall death rates for some broad categories such as heart disease or malignant neoplasms changed little overall, the death rates by individual cause did change. For example, the rate for cardiac dysrhythmias declined 38.5 per cent, from 15.6 to 9.6 per 100,000 population, and malignant neoplasms of an unspecified site fell 28.1 per cent, from 8.9 to 6.4. At the same time, the rate for primary liver cancer increased from 0.8 to 1.7, a 112.5 per cent rise. Although the overall death rate for accidents fell 10.3 per cent, the rate for falls increased 8.9 per cent, from 5.6 to 6.1. Some other changes include: a 43.8 per cent decrease in deaths due to aspiration, a 46.2 per cent decrease in deaths attributed to an unspecified intestinal obstruction, a 90.9 per cent decrease in deaths due to motor vehicle traffic accidents of unspecified nature, a 16.7 per cent increase in deaths due to urinary tract infections, and a 30.3 per cent increase in deaths due to suicide with a specified firearm.

TABLE 1—Percentage Difference in the Death Rates of the 20 Leading Causes of Deaths before and after Querying, Oregon, August 1986–July 1987

Cause of Death	ICD Codes	Rate* Before	Rate* After	Per Cent Difference
Peptic Ulcer	531–533	2.3	3.0	+30.4
Alcoholic Disease	291, 303.0, 357.5, 425.5, 535.3, 571.0–571.3	9.4	11.8	+25.5
Hypertension	401, 403	3.7	4.2	+13.5
Other Diseases of Arteries	441–448	9.4	10.0	+6.4
Chronic Obstructive Pulmonary Diseases	490–496	42.6	44.7	+4.9
Diabetes Mellitus	250	12.6	13.2	+4.8
Arteriosclerosis	440	15.2	15.8	+3.9
Homicide	960–969	5.3	5.5	+3.8
Congenital Anomalies	740–759	5.3	5.5	+3.8
Malignant Neoplasms	140–208	198.8	202.8	+2.0
Early Infancy	760–779	5.1	5.2	+2.0
Cerebrovascular Disease	430–438	72.9	74.3	+1.9
Suicide	950–959	16.4	16.4	0
SIDS	798.0	4.2	4.2	0
Undetermined External Causes	980–989	1.6	1.6	0
Heart Disease	390–398, 402, 404.0–425.4, 425.6–429.9	296.6	295.9	−0.2
Influenza and Pneumonia	480–487	28.5	28.4	−0.4
Accidents	800–949	48.7	43.7	−10.3
Hernias and Intestinal Obstruction	550–553, 560	2.3	2.0	−13.0
Nephritis and Nephrosis	580–589	6.2	4.2	−32.3

*Rate per 100,000 population.

Because the OCHS has had a query program for many years, and consequently many physicians have learned how to complete a death certificate thoroughly and accurately, the effect demonstrated in Table 1 is very conservative. One measure of how quickly physicians learn to provide the necessary information is seen with the recently introduced tobacco use query. The OCHS initially requested that physicians note on death certificates the role of tobacco, if any, in the deaths of their patients. The request and rationale for this was presented in an article in the Oregon State Health Division's biweekly publication that is sent to all Oregon physicians. In the first three months, 60.0 per cent of the lung disease deaths were queried because the certifiers failed to note the role of tobacco. During the second quarter, the figure declined to 55.0 per cent, and by the third quarter only 34.0 per cent of the deaths required querying.

Survey of State Querying Practices

In the survey of other state vital statistics programs, twenty-one (40.4 per cent) of the registrars reported that their nosologists followed NCHS guidelines specifying the priority levels used. Four states responded they used the guidelines but gave vague answers as to the priority levels used. The four states (and their responses) are: Kansas (depends); Michigan (basic); New Jersey (as needed); Tennessee (partial). At least 23 states do not query at the minimum priority level, despite the fact that the NCHS has stated, "It is expected that every registration area will fulfill, as a minimum, all of the priority Level 1 requirements."[1] Some states, such as Rhode Island and Mississippi, did not report following NCHS guidelines, but appear to have thorough query programs and may be meeting these requirements.

Among those states following the NCHS guidelines, priority Level 1 was most often used (Figure 1). While registration areas using priority Levels 1–3 meet the NCHS minimum standard, Level 4 provides more comprehensive information. Levels 5 and 6 are "not necessary for general querying."[1] Oregon queries at Level 4, while four states—Colorado, Illinois, Kentucky, and South Carolina—reported using priority Level 5, the highest consistently used level. Colorado and Kentucky also occasionally use Level 6. Five states, then, reported querying at a comprehensive level. However, follow-up phone calls to states reported to query at Level 5 revealed that some do not query all the items within priority Level 5 or do so depending on workload.

Discussion

A query program can make a dramatic difference in the accuracy of cause-of-death data. Researchers comparing geographic differences in mortality should be aware of cause-of-death query programs and how these programs ultimately affect the data. For example, in 1984, the reported maternal death rate for the United States was 6.0 deaths per 100,000 live births.[5] The Oregon maternal death rate (13.6) during 1982–86 was 126.7 per cent higher than the national rate. These data suggest that a rigorous query program can substantially increase the reporting of deaths attributed to complications of pregnancy, childbirth, and the puerperium.

A computerized query tracking system allows the usefulness of responses to be monitored. Queries may be "fine-tuned" in this way to elicit better quality information and to measure the impact of rewording queries. Sometimes a change does not yield the expected improvement. For example, during August 1986 through January 1987 cardiac arrest had been queried with Form 1, "Was the underlying cause of the cardiac arrest known?" Over one-half (55.7 per cent) of 122 queries yielded information that resulted in a more specific underlying cause-of-death. In February, the query was reworded to "Some terminal events, such as cardiac or cardiopulmonary arrest, seldom arise spontaneously: Was there an underlying or more serious condition causing the cardiac or cardiopulmonary arrest?" This new phrasing resulted in only 37.3 per cent of the 107 queries with more specific causes of death. Additionally, 13.1 per cent of the physicians failed to respond to the new phrasing com-

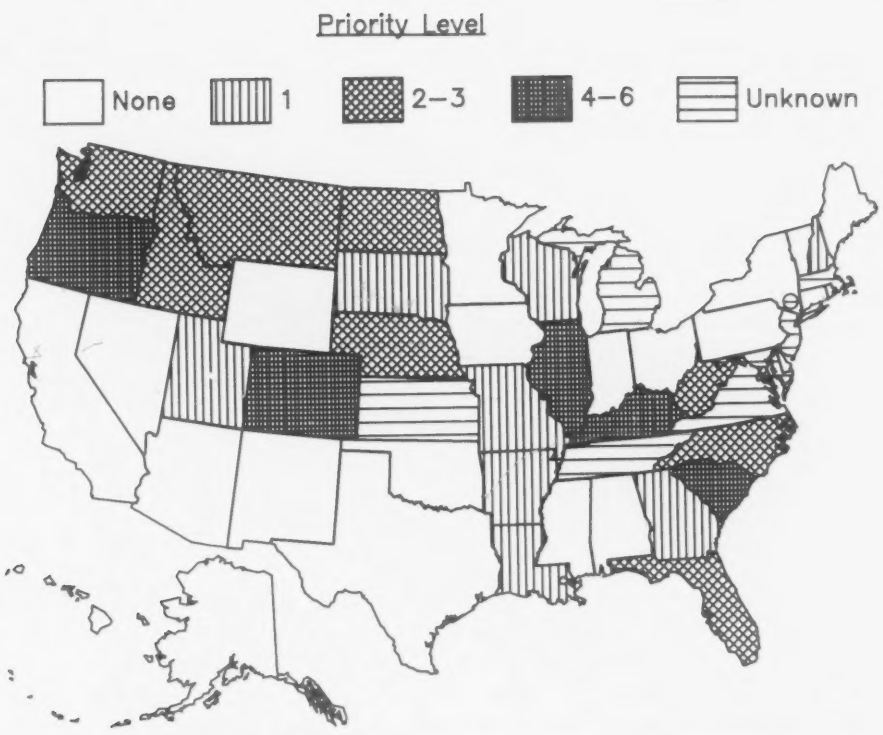

FIGURE 1—Survey Results of States' Querying Practices in Relation to NCHS Guidelines and the Priority Levels Used

pared to 5.7 per cent previously. Consequently, the previous form was reinstituted for this cause.

Unspecified liver disease and cirrhosis are priority Level 3 items under the NCHS guidelines. Data from states where death certificates are not queried at a level equivalent at least to priority Level 3 will not have as complete alcohol-related mortality data as areas which do query at this level. Only 10 states indicated on the survey forms that they consistently queried at least at priority Level 3 or for unspecified liver cirrhosis and disease. Four of these—Florida, Idaho, Kentucky, and Washington—during follow-up phone calls in June and July 1987, stated that they did not query for these causes, although they had listed Levels 3, 3, 5–6, and 3, respectively, on the survey form. Of the remaining six states, only three queried prior to the publication of NCHS guidelines to determine if cirrhosis or liver disease was alcohol-induced. The survey by phone included the question, "When did you begin to query liver cirrhosis and disease deaths to determine the role of alcohol?" Responses given were: Colorado, 1985; Illinois, 1985; Nebraska, at least 20 years ago; North Dakota, early 1970s; South Carolina, two to three years ago.

Because the OCHS has queried physicians about unspecified cirrhosis or disease of the liver since the early 1970s, many physicians report alcoholic or Laennec's cirrhosis without the necessity of a query. During 1986, 294 certificates were coded to 571.0–571.9. Most did not require querying because the decedent's liver disease or cirrhosis was noted to be alcoholic in origin (200) or because another cause was listed (22). Fifty-four certificates with ICD codes 571.5 and 571.8 were queried and all but nine (83 per cent) reassigned to 571.0–571.3. An additional 11 deaths were discovered to be due to alcoholic liver disease and cirrhosis from queries seeking unrelated information, while seven others were assigned to 571.4 and 571.8 for the same reason.

Oregon's query program has resulted in the state reporting the highest percentage of liver cirrhosis and disease deaths attributed to alcohol abuse. Nationally, 41.7 per cent of all liver disease and cirrhosis deaths in 1984 were listed as due to alcohol.** In Oregon 82.4 per cent were so classed, similar to the 83 and 85 per cent reported by Garceau.[6,7] Nonetheless, the overall liver cirrhosis and disease death rate was only marginally higher in Oregon than the US, 12.0 and 11.6 per 100,000 population, respectively. The apparent per capita alcohol consumption*** in Oregon (2.63) closely mirrors the national average (2.65).[8]

States whose vital statistics programs query at priority Level 4 or higher can be expected to have improved mortality data not only for alcohol-related deaths, but for all causes of death. While querying physicians in order to code and classify causes of death in accordance with ICD concepts, rubrics, and codes is just one step in the collection and dissemination of public health statistics, it is one that can, in certain instances, dramatically affect mortality data. Varia-

**Unpublished data. Division of Vital Statistics, National Center for Health Statistics, Hyattsville, MD.

***Gallons of absolute alcohol apparently consumed based on sales/revenues collected by retailers, divided by the population aged 14 and over.

tion in cause-of-death querying between registration areas is one more factor that points to the need for a national program to periodically assess the quality of mortality data.[9] For mortality statistics to be of value as a national surveillance mechanism, it is important that data be collected in a standard and consistent manner. The NCHS recommendations, published only recently, may be adopted by more states as the benefits from establishing a query program are documented.

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APPENDIX A
Causes of Death Suggestive of AIDS and Queried When AIDS is Not Mentioned on the Death Certificate

Atypical mycobacteria (*Mycobacterium avium* or *intracellulare*)
Bronchial, pulmonary or esophageal candidiasis
Chronic lymphoid interstitial pneumonitis
Cocciodioidomycosis
Cryptococcal meningitis
Cryptosporidiosis
Cytomegalovirus infection
Disseminated histoplasmosis
Hemophilia
Herpes virus infection
Immune deficiency
Immunoblastic sarcoma or lymphoma
Isosporiasis
Kaposi's sarcoma
Lymphoma limited to the brain
Non-Hodgkin's lymphoma in a person under age 60
Pneumocystic pneumonia
Pneumonia in a person between the ages of 1 to 50 and no clear alternative to AIDS as an explanation for the person's unusual susceptibility to pneumonia (i.e., leukemia, splenectomy)
Progressive multifocal leukocephalopathy
Salmonella (non-typhoid)
Small non-cleaved lymphoma (Burkitts' tumor or Burkitt-like lymphoma)
Stronglyloidosis
Toxoplasmosis
Tuberculosis in a person under age 60

APPENDIX B
Causes of Death Queried in Addition to Priority Level 4 by the Oregon Center for Health Statistics

Cause	ICD Code
Septicemia	038
Serum hepatitis	70.2–.3
Malnutrition	263
Dehydration and electrolyte disorders	276.5, .7–.9
Aplastic anemia	284
Hydrocephalus not specified as congenital (or with a duration since birth)	331.3–.4
Pulmonary embolism among persons under age 65 years	415.1
Pulmonary heart diseases (except kyphoscoliotic heart disease)	416.9
Varicose veins of esophagus	456.0–.1
Pneumonia in categories 481, 482, 485 or 486 with mention in Part II of carcinoma	481, 482, 485, 486, 507.0
Any pneumonia stated to be terminal	480–486, 507.0
Rupture of the esophagus	530.4
Intestinal obstruction with or without mention of adhesions	560.8–.9
Aspiration not indicated as accidental or due to a disease condition	911, 912
Liver disease not specified as alcoholic when *not* selected as the underlying cause of death	571.5, .7–.9
Type of gun used in intentional and unintentional deaths	922.9, 955.4, 985.4
Passenger status in selected transportation accidents	800–839, .9s only
Hemorrhage of any site except cerebral or related to childbirth	
Any external cause of death with items 36a–36f not completed	
Females in the following age and cause of death categories for maternal deaths	*Ages 17–35* 303 571.0–571.3 *Ages 17–39* 001.0–139.9 240.0–302.9 304.0–337.9 341.0–358.9 360.0–429.9 438–570.9 571.4–796.9 *Ages 17–44* 430–437.9 797.0–799.9

Two Variants of Coronary Atherosclerosis in Six Populations

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Abstract: Coronary arteries and aortas obtained at autopsy in six populations were examined microscopically. Two variants of atherosclerosis could be separately quantified by objective morphometry. These two variants, fibroplasia and atheronecrosis, were both found to increase with age and to correlate with each other from one autopsy to the next after age adjustment. Despite the correlation of these two variables among individuals, the measures of fibroplasia and atheronecrosis did *not* correlate with each other across population groupings. A positive correlation of this kind between some populations was canceled by an inverse correlation between other populations. For example, Manila was the population with the greatest intimal fibroplasia in the coronary arteries at each age, and Sao Paulo was among those with the least intimal fibroplasia. Nevertheless, Sao Paulo showed the greatest tendency toward atheronecrosis, while Manila was among the populations with the least tendency toward atheronecrosis. The possibility arises, therefore, that the subjects from Sao Paulo were exposed to a different assortment of etiologic agents than were the subjects from Manila, and that their arteries, therefore, manifested a different kind of atherosclerosis. (*Am J Public Health* 1989; 79:575–581.)

Introduction

Atherosclerosis progresses throughout life from childhood precursors into eventually fatal complications in late life that affect coronary, cerebral, and other arteries. The rate of progression of atherosclerosis is thought to vary among human populations. The major basis for this belief is the International Atherosclerosis Project (IAP).[1] Vital statistics, clinical observations, and spotty reports of autopsy findings from nation to nation have consistently ranked some countries as high and others as low in the incidence and prevalence of atherosclerotic complications. An exact comparison of this kind, however, requires that representative samples of arteries from various sources be brought together in one place and objectively evaluated. This has been done twice: the World Health Organization (WHO)-sponsored five towns study,[2] and the IAP.[3] The WHO study found little difference among the five towns that they chose to compare; the least average percentage of surface affected by raised lesions being 30.6 in Yalta and the highest being 45.4 in Prague (standardized to age 55). The IAP compared arteries from populations which represent the extremes of known human experience; ranges were from 6 per cent in the Durban Bantus to 24 per cent in the New Orleans Whites (standardized to age 40).

The principal method used for measuring atherosclerosis in the IAP and in the WHO study was to assess by the unaided eye the extent of intimal surface judged to be affected by atherosclerotic lesions. This method was sufficient to show that: 1) raised lesions occur in the aortas and coronary arteries of all human populations, and 2) variations among populations are relatively small in comparison with variations among individuals within populations and between age groups. Therefore, the epidemiology of atherosclerosis requires rather large sample sizes to demonstrate the modest differences that prevail between populations. This method of gross inspection, however, was unable to probe the microscopic structure of aging arteries.

Two broad classes of raised atherosclerotic lesions are recognized microscopically, those with and those without necrotic cores.[4–6] These have been called "atheroma" and "preatheroma".[5] Microscopic examinations of aortas from the IAP archive have led to the suggestion that the pathogeneses and etiologies of these two classes of raised lesions are not the same in all populations.[7] Some populations such as Bogota residents and Durban Bantus were low on measures for both types of lesion while other populations, such as the Durban Indian, were high on both. On the other hand, the Sao Paulo, Brazil population, with the *greatest* tendency toward aortic lesions of the atheronecrotic type, was also among the populations with the least tendency toward the nonnecrotic, "fibroplastic" type. The population with the *least* tendency to atheronecrosis in that study of aortas, New Orleans Blacks, was also among the populations with the greatest tendency to fibroplasia. Those results raised the possibility that two separate processes, caused by different sets of etiologic agents, might be interacting in the arterial intima.

Microscopy is a little-used technology that offers unexplored new opportunities for seeking understanding of atherosclerosis epidemiology. The previous findings from the microscopic examination of aortas in eight populations are extended for this report to the coronary arteries in six populations.

Methods

Selection of Cases

The IAP, conducted in 1960-64, brought together in New Orleans about 23,000 sets of coronary arteries and aortas from 14 locations, five of them biracial, 19 location-race groups altogether.[1] In a previous study, aortas from eight of these populations were examined microscopically.[7] The eight groups chosen from the collection represent the extremes of observed arterial involvement by raised lesions and fatty streaks. Coronary arteries were available from the same cases which provided aortas in six of these groups: Durban, South Africa (Bantu and Indian); Bogota, Colombia; Mexico City, Mexico; Manila, Philippine Islands; and Sao Paulo, Brazil (Whites only).

A total of 395 specimens were successfully retrieved and prepared for microscopy, most of the failures being due to drying of the formalin used for storage. Numbers of cases by age, sex, population, and cause of death are given in the various tables.

Cause of Death Categories

Cases which clinically or at autopsy showed some atherosclerosis-related feature were classed as "Athero", all others

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were called "Basal". Atherosclerosis-related features include hypertension, diabetes, myocardial scar or cerebral scar when observed incidentally, as well as overt causes of death such as myocardial infarct, stroke, or ruptured aneurysm.[10]

Processing of Tissues

Aortas and coronary arteries were previously opened longitudinally, formalin fixed, sudan stained, sealed in formalin-filled plastic bags, and stored in cardboard boxes containing stacks of 100 sets. Aortas were newly sampled along the left and right lateral walls from the 4th to the 12th intercostal ostia, as detailed elsewhere.[7] Right, left circumflex, and anterior descending coronary arteries were each cut into four segments of 2 cm length, some segments being absent in vessels shorter than 8 cm.[8] Longitudinally oriented segments were decalcified with acetic acid, blocked in paraffin, sectioned at 11 μm, and stained with hematoxylin-eosin. Ink marks were placed on the cover slips to define one centimeter intervals on the coronaries, and at intervals in the aorta to define 1/20 of the sample length—approximately 0.7 cm in most aortas. Up to 27 marked positions were evaluated in the coronaries and up to 40 in the aortas.

Evaluation of Slides

Positions affected by atheronecrosis, operationally defined as having cholesterol clefts easily discerned under the 40x objective lens, were marked with black ink; the percentage of the length of the specimen marked positive for atheronecrosis was called P_{Aa} for the aorta and P_{Ac} for the coronaries. The mean thickness of the intima (μm), averaged over the nonnecrotic points, were called F_c for coronaries and F_a for aortas. Details of these methods are discussed elsewhere.[8]

Statistical Strategy

The outcome measures for the two types of atherosclerotic lesions in the two arterial segments, P_{Aa}, P_{Ac}, F_a, and F_c, were found to correlate with age and with each other from one case to another. Variation of these measures among populations was therefore examined after age adjustment. To test if the age adjustment was equally appropriate to all populations, the adjusted values of F were compared separately for the ages over and under 30 years. This could not be done for P_A, because this measure is nearly always zero below age 30. Next, the residual variation of P_A among populations was examined after adjustment for its correlation with F; this adjustment was improved by taking the square root of P_A.[9] The analyses were done separately for two broad cause of death groups to examine the effect of inequalities among the autopsy samples in their cause of death compositions. The statistical details of the analyses are reported in the Appendix.

Results

The average nonnecrotic intimal thickness, i.e., *fibroplasia*, tended to rise with age, as shown in Figure 1. Perhaps the most striking feature of the scatter plot is the great variation among individuals who are alike in age. However, the average trend, as shown by the curve representing equation (1) is clearly upward with age.

Atheronecrosis tended to affect the older cases and those with the greatest intimal fibroplasia, irrespective of cause of death category. In Figure 1, the open circles, representing specimens *with* atheronecrosis, are upward and rightward of the closed circles representing specimens *without* atheronecrosis. The dashed curve in Figure 1 represents the optimal separation between the open and closed circles, as given by equation (2). The curve helps to emphasize the finding that

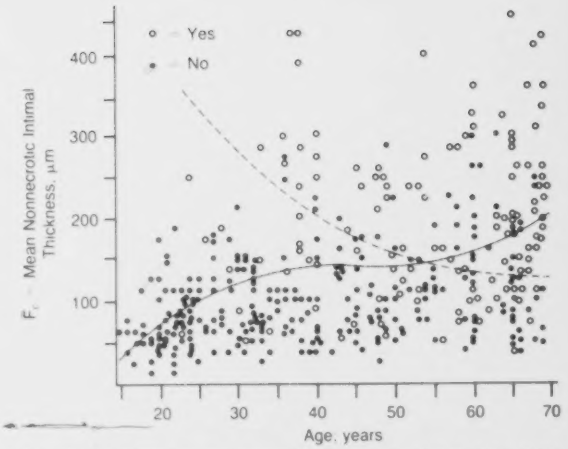

FIGURE 1—All 395 cases are represented, using open symbols to designate the 132 cases with at least one instance of atheronecrosis in the coronary arteries. The solid curve represents the cubic regression equation (1). The dashed curve represents the logistic regression equation (2) with log odds = 0, which optimally separates the open from the closed circles.

elderly subjects can manifest atheronecrosis at lesser degrees of fibroplasia than is required in youth.

Variation of Fibroplasia among Populations

Some of the variation among individuals seen in Figure 1 is due to population membership, because populations were not all alike in fibroplasia. Manila was the population with the greatest intimal thicknesses at each age in the coronary arteries and Sao Paulo was among those with the least. As shown in the scatter plot of Figure 2, 58 of the 74 specimens of coronary arteries from Manila are above the age-adjustment regression line. From Sao Paulo, 12 of the 69 specimens are above the line.

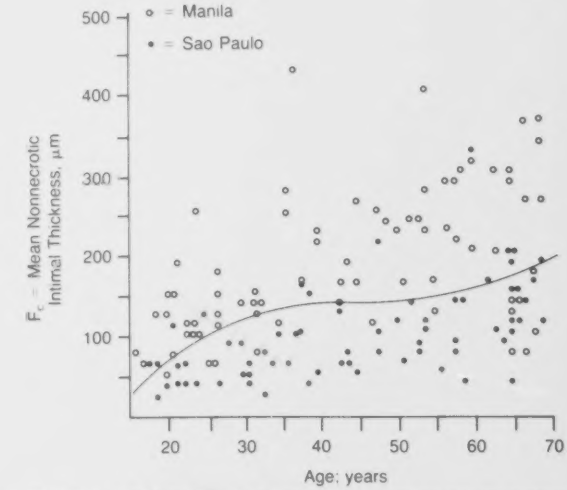

FIGURE 2—Manila and Sao Paulo illustrate the populations with the extremes of high and low fibroplasia observed in the intima of coronary arteries at each age. The curve represents the cubic regression equation (1).

TABLE 1—Age-Adjusted Mean Nonnecrotic Intimal Thickness* (μm) by Arterial Segment and Cause of Death for Six Populations

Populations	Coronaries by Cause			Aortas by Cause			Numbers of Cases	
	Total	Basal	Athero	Total	Basal	Athero	Basal	Athero
Manila	189	172	205	308	266	350	37	37
Durban Indian	148	130	176	279	274	287	43	27
Bogota	139	142	135	234	209	284	42	21
Mexico	129	123	157	198	201	184	51	12
Sao Paulo	97	94	103	213	194	261	50	19
Durban Bantu	95	87	123	198	182	259	45	11
Total	135	123	161	242	218	292	268	127
ANOVA**		F (P-Value)			F (P-Value)			
Causes		24.98	(0.00)		32.01	(0.00)		
Populations		10.26	(0.00)		5.59	(0.00)		
Interactions		1.26	(0.27)		1.40	(0.21)		

*Adjusted to age 45 through the regression equations (1) and (1a).
**Results of 2-Way analysis of variance, F-Value (Pr>F), are from PROC GLM of the SAS package carried out upon the residuals from the cubic regression equations.

The population differences persisted after age adjustment (Table 1) by regression equations explained in the Appendix. This analysis retains the distinction between cause of death groupings to show that atherosclerosis-related causes did not greatly distort the pattern of variation among populations as seen in the basal cases only. Cause and population differences were present for the coronary arteries as well as the aortas. Manila was consistently ranked first or second, and Sao Paulo fourth, fifth, or sixth for aortas and coronary arteries of both cause of death groupings.

In Figure 2, the specimens from Manila seem to show greater fibroplasia than those from Sao Paulo even at ages less than 30 years. This result was confirmed by ANOVA in Table 2. Intimal fibroplasia of the coronary arteries in the six populations before and after age 30 showed a rank correlation of 0.99. The comparable correlation for the aortas is 0.72. The age-adjusted fibroplasia analyzed in Table 2 showed no differences between groups over and under age 30, suggesting that the age adjustment is equally appropriate to all six populations.

Variation of Atheronecrosis among Populations

The six populations differed from each other in raw measures of the percentage of atheronecrosis in the coronary arteries (Figure 3). As seen in the diagram, the differences among populations in atheronecrosis seem independent of the variation in fibroplasia. We have already seen, however, that specimens with more fibroplasia are, for each age, at greater risk of showing atheronecrosis (Figure 1, equations (2), (3), (3a)). It seems fair to ask, therefore, if the amounts of atheronecrosis observed in each population might be commensurate with the degrees of fibroplasia, after appropriate age adjustment. Manila, for example, had the greatest fibroplasia, while Sao Paulo had nearly the least; therefore, according to equation (3), these populations should be respectively among the greatest and the least affected by atheronecrosis. The reverse outcome was actually observed, as given in Table 3. Sao Paulo had the *most* atheronecrosis, after adjustment for age and fibroplasia and Manila was among the populations with the least atheronecrosis (adjusted); the respective means ± SEMs were 1.97 ± 0.22 and 0.70 ± 0.18 square rooted percentage units. In a graphic presentation of these results, Figure 4, the line representing equation (3) is plotted, and departures from that line represent the residuals used in the ANOVA of Table 3. The two populations had equal numbers of non-zero specimens, 32 out of 74 for Manila and 32 out of 69 for Sao Paulo, despite the fact that the risk functions from equations (3) and (3a) predict greater risk in Manila. In Figure 4, 30 of the 32 non-zero specimens from Sao Paulo have values for atheronecrosis

TABLE 2—Age-Adjusted Mean Nonnecrotic Intimal Thickness* (μm) by Arterial Segment and Age for Six Populations

Population	Coronaries by Age (yrs)			Aortas by Age (yrs)			Numbers of Cases by Age (yrs)	
	Total	<30	30–69	Total	<30	30–69	<30	30–69
Manila	189	164	199	308	269	324	21	53
Durban Indian	148	142	150	279	253	290	20	50
Bogota	139	138	140	234	217	237	10	53
Mexico	129	131	128	198	221	190	15	48
Sao Paulo	97	112	112	213	219	211	13	56
Durban Bantu	94	112	89	198	209	194	13	43
Total	135	136	135	242	235	244	92	303
ANOVA**		F (P-Value)			F (–P-Value)			
Ages		0.06	(0.81)		0.39	(0.53)		
Populations		12.14	(0.00)		7.58	(0.00)		
Interactions		1.20	(0.31)		0.74	(0.62)		

**Age adjustments and ANOVA as in Table 1. The purpose of doing this analysis on the age-adjusted variable is to test if the same age adjustment is appropriate to all six populations.

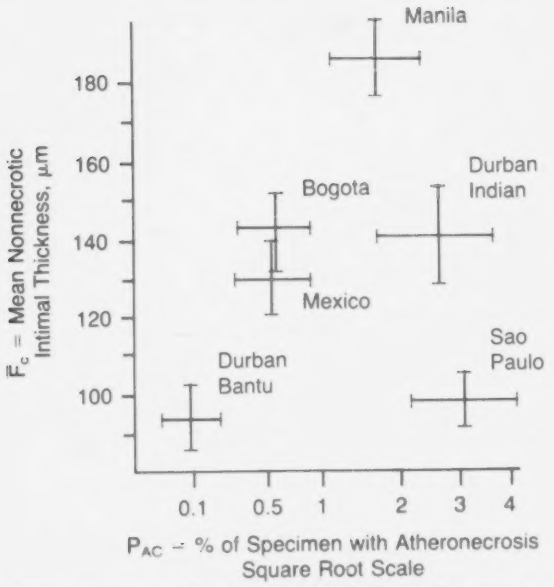

FIGURE 3—The means, plus or minus one standard error of the mean, are represented for each of the 6 populations. Fibroplasia and atheronecrosis values plotted here are not age-adjusted or otherwise adjusted.

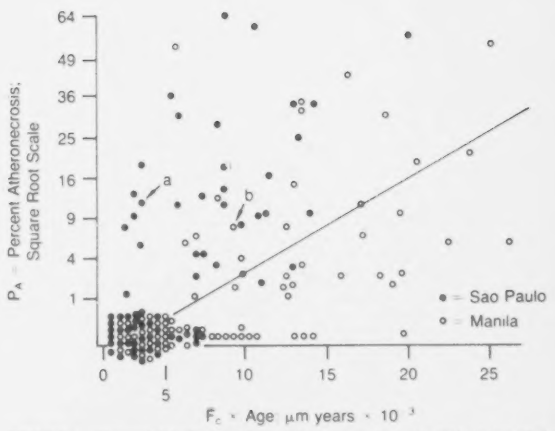

FIGURE 4—The variable obtained by multiplying age times fibroplasia is plotted against atheronecrosis, on a square root scale, for the two populations which differed most on plots of this kind. The product variable, fibroplasia-years, was especially interesting because of its importance in equations (3) and (3a). The cases marked "a" and "b" were near the centers of the scatters for *young* persons in the two populations, and were therefore considered to be "typical"; they are illustrated photographically in Figure 5.

that are above the regression line, compared with 15 of those from Manila. The six populations were not ranked the same on measures of fibroplasia and atheronecrosis (adjusted) (rank r = −0.03).

Photographic Examples

Two cases were chosen to illustrate photographically the contrast between typical appearances in Sao Paulo and Manila in Figure 5. Case "a" from Sao Paulo had age = 34, F_c = 62, and P_{Ac} = 9.4. Case "b" from Manila had age = 44, F_c = 192, and P_{Ac} = 8.0. These two cases are typical because of having values that are near the average for subjects in their respective populations. The photographs illustrate that nonnecrotic intimal thickness is less in Sao Paulo than in Manila, yet atheronecrosis can unexpectedly exist in this less thickened fibroplastic intima.

Attrition of the Sample

Specimens from 10 males and 10 females were drawn from the archive in each of the six 10-year age groups, making 120 for each of the six populations. Forty-five per cent of the specimens were found to be unsuitable for histology; these had been improperly dissected, fixed, or stored, the major cause of attrition being drying during storage. By Chi-Square test, the attrition of specimens appeared to be randomly distributed among the six populations. However, the population with the thickest specimens, Manila, had the least attrition (46) while the population with the thinnest specimens, Durban Bantu, had the most attrition (64), and the rank correlation in this regard was statistically significant (r = 0.829, Table 4). It seems likely that the thinnest specimens are most prone to being ruined by improper storage. How this

TABLE 3—Mean Atheronecrosis, Adjusted for Age and Nonnecrotic Intimal Thickness,* Expressed as Square Root of Percentage of Specimen, by Arterial Segment and Cause of Death for Six Populations

	Coronaries by Cause			Aortas by Cause			Numbers of Cases	
Populations	Total	Basal	Athero	Total	Basal	Athero	Basal	Athero
Manila	0.70	0.59	0.81	1.07	0.55	1.58	37	37
Durban Indian	1.54	1.21	2.06	0.94	0.79	1.18	43	27
Bogota	0.49	0.17	1.13	0.39	0.34	0.51	42	21
Mexico	0.72	0.65	0.99	0.44	0.45	0.39	51	12
Sao Paulo	1.97	1.60	2.95	1.13	0.98	1.51	50	19
Durban Bantu	0.65	0.51	1.23	0.27	0.33	0.02	45	11
Total	1.03	0.82	1.47	0.75	0.61	1.05	268	127
ANOVA**		F (P-Value)			F (P-Value)			
Causes		18.07	(0.00)		11.66	(0.00)		
Populations		8.94	(0.00)		4.84	(0.00)		
Interactions		2.33	(0.03)		4.71	(0.00)		

*Adjusted to age 45 and to intimal thickness of 135 μm (Coronaries) or 240 μm (Aortas) through the regression equations (3) and (3a)

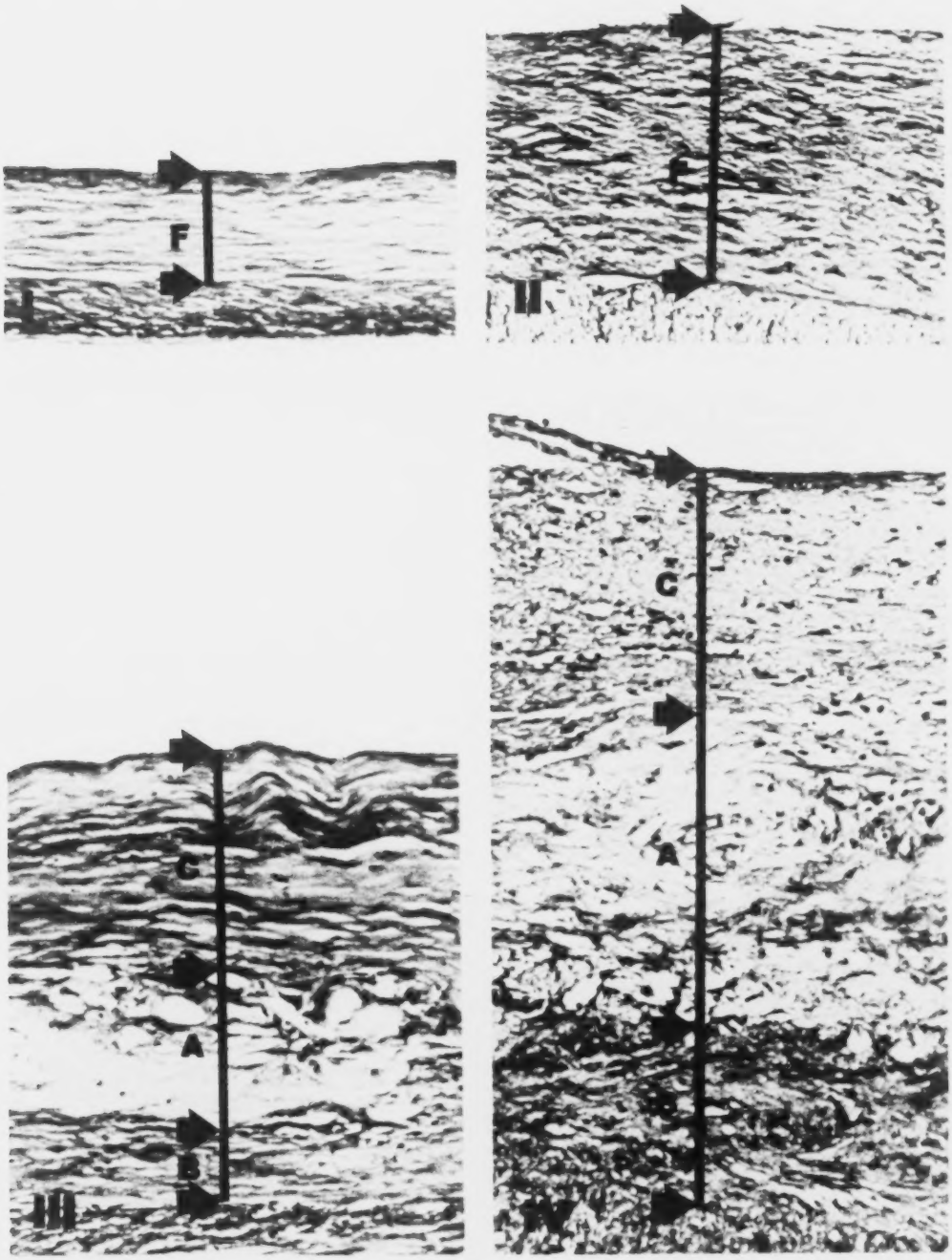

FIGURE 5—Coronary arteries of the two cases marked "a" and "b" in Figure 4 are illustrated in frames I and III and frames II and IV, respectively. Frame I shows nonnecrotic intimal thickness of 80 μm (F) which was near the average for this case and for Sao Paulo males of age 45. Frame II shows thickness of 185 μm which is near average for this case and for Manila males of age 45. Frames III and IV show necrotic cores (A) accompanied by "threshold" amounts of fibroplasia formed from the sum of the fibrous cap (C) and base (B). B + C = 200 μm in the Sao Paulo case of Frame III and 300 μm in the Manila case of Frame IV, amounts which are near average for 45 year old males in those two populations. Hematoxylin-eosin, X280.

selection bias might affect the results is hard to assess, but is apt to be small.

Discussion

Atheronecrosis is generally believed to be a late complication that sometimes intrudes upon fibroplastic, "preatheromatous", intimal thickening. This viewpoint might look upon equation (3) not just as an empirical description of an observed correlation, but as a model of a real process evolving in time. This viewpoint was recently elaborated in some detail.[8] It is not necessary, however, to accept this theory in order to pursue the epidemiology of the tissue abnormalities. Manila was observed in the data reported here to have a greater tendency toward fibroplasia than Sao Paulo but a lesser tendency for the thickened and aged intima to show atheronecrosis. Since "the cause" of atherosclerosis is multifactorial, the subjects from Sao Paulo may have been exposed to a different assortment of etiologic agents than were the subjects from Manila, and therefore their arteries manifested a different kind of atherosclerosis.

A potential selection bias which could, in theory, bear upon the observed outcome is the so called "Berkson fallacy."[10] If two populations differ in death rates or autopsy rates for certain kinds of causes, such as malnutrition, infections, etc., then the two series of autopsies could reveal spurious differences in atherosclerosis; this is because subjects with coronary heart disease might comprise most of the autopsies in one place but not the other. This possibility seems an unlikely explanation for the differences between populations reported here, because the same pattern of population contrasts was seen in both of the cause-specific groups of autopsies. Manila had the greatest fibroplasia in both basal and atherosclerosis-related cause of death groups, while Sao Paulo ranked fifth or sixth in both cause groups (Table 1). The two populations reversed their relative positions when looking at atheronecrosis instead of fibroplasia, and this reversal was seen equally well in both cause of death groups (Table 3). Indeed, the contrasts between populations seem somewhat greater in the atherosclerosis-related than in the basal groups, as if death from one kind of atherosclerosis might cut short the progression of the other kind. The same result was previously seen in the aortas.[7]

Another kind of selection bias must also be considered. These speculations do not necessarily apply to the general populations in Brazil or the Philippines, or elsewhere. To what extent the particular racial and cultural group samples studied here might be representative of the source cities or countries constitute questions that are newly raised by the findings. Without knowing whether these cases were sampled from populations which are representative of their nations of origin, it is still possible to conclude that fibroplasia and atheronecrosis do not always travel together across population boundaries. This is a newly discovered principle, previously seen in the aortas, but, for the coronary arteries reported here for the first time. The matter requires further verification before general acceptance.

A dissociation between fibroplasia and atheronecrosis has been noted in some experimental situations. For instance, rabbits showed fibroplastic intimal responses to feeding peanut and coconut oils, atheromatous reactions to butter fat, and little of either result from corn oil.[11] It could be no more than coincidence that the arteries from a land of coconut oil consumption, Manila,[12] had a resemblance to those of rabbits fed oil of coconuts, or that arteries from a land of cottonseed and soya oil consumption, Sao Paulo,[12] resemble those of rabbits fed butter. The experimental data, however, do help to validate the suggestion found in these human data that atherosclerosis may be a composite result of two separate processes interacting in the same tissue. This matter has been raised into serious consideration by the findings reported here, because comparisons of arteries from different human populations showed an effect similar to that previously known only for experimental settings. The next step, which is to correlate the two kinds of atherosclerosis with putative etiologic agents, will require further comparisons of human populations.

APPENDIX

Quantitative variables measured upon two cause of death classes in six populations lend themselves to multivariate statistical analysis. The SAS package of programs, version 5, was used.

Fibroplasia in the coronary arteries of 395 cases were related to age by a cubic regression equation with all coefficients statistically significant

$$F_c = 0.003{,}12\ A^3 - 0.417\ A^2 + 19.2\ A - 168 \quad (R^2 = 0.154), \qquad (1)$$

where F_c is fibroplasia in the coronary arteries, measured in micrometers, averaged over all of the points observed in a specimen, and A is age in years. All coefficients are significant at $\alpha = 0.05$. The curve representing this equation is drawn in Figures 1 and 2. The linear correlation, with $R^2 = 0.142$, is not greatly inferior to the cubic, and might suffice for age adjustment. Nevertheless, differences in age structure among the populations are apt to be better compensated by the cubic, and the equation (1) is therefore used.

For the aortas, the higher order terms in the cubic regression were not statistically significant ($\alpha = 0.1$), and were therefore not retained. For age adjustment of fibroplasia measures in the aorta (F_a),

$$F_a = 5.34\ A + 1.45 \quad (R^2 = 0.304) \qquad (1a)$$

Residuals from equations (1) and (1a) were examined by ANOVA in Tables 1 and 2.

Atheronecrosis can be treated as a quantitative variable, expressed as the percentage of affected specimen (P_{Ac} and P_{Aa} for coronary artery and aorta respectively), or as a categorical variable, each specimen treated as having or not having at least one instance of atheronecrosis. Setting odds equal to the ratio of affected to not affected specimens, logistic regression yielded

$$\text{Log odds} = 0.000{,}304\ F_c \cdot A - 2.80 \quad (\chi^2 = 103.9) \qquad (2)$$

This function correctly classifies 79.3 per cent of cases, and the χ^2 of 103.9 refers to the 2×2 classification table. This equation, with log odds = 0, is plotted in Figure 1. The linear terms in F_c and A were rejected as not statistically significant. By multiple linear regression

$$\sqrt{P_{Ac}} = 0.000{,}211\ F_c \cdot A - 0.254 \quad (R^2 = 0.354) \qquad (3)$$

Again, the linear terms were not significant. For the aorta, the linear term on age was retained because it was significant,

$$\sqrt{P_{Aa}} = 0.000{,}164\ F_a \cdot A - 0.030{,}3A + 0.35 \quad (R^2 = 0.542) \qquad (3a)$$

Residuals from equations (3) and (3a) are examined by ANOVA in Table 3.

ACKNOWLEDGMENTS

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Survey Shows HIV Reporting Practices in 50 States

A national survey of HIV reporting requirements in all 50 states and the District of Columbia—conducted by the AIDS Policy Center of the Intergovernmental Health Policy Project at George Washington University—shows that anonymous testing, in some form, is available in more than four-fifths of the states. The results demonstrate a clear trend to maintain public access to anonymous testing while ensuring that health officials have access to test recipients' names under certain conditions.

In most states, the system reflects the efforts of public health officials to balance various needs, including privacy and confidentiality, risk-reduction education, follow-up care, intervention, and monitoring spread of the disease.

According to the survey, state reporting systems fall into five major categories based on whether those who test positive for HIV must be reported to state or local health officials. The categories are: 1) confidential reporting (where names must be reported); 2) confidential reporting, with some anonymous testing available; 3) anonymous testing with names required in specific instances; 4) anonymous reporting, where aggregate or demographic data must be reported or where individuals' test results must be reported, without names or identifiers; and 5) anonymous testing where no names or individual information is reported.

Findings of the survey indicate that:

- only eight jurisdictions require names to be reported as a matter of policy or law;
- 10 states have retained their anonymous test sites without mandating other reporting requirements;
- a dozen states either require aggregate reporting of seropositive patients (without names) or allow codes to be used in lieu of names;
- six states have determined that HIV reporting should remain basically anonymous with a few mandated exceptions, such as name reporting of seropositive infants and persons who indicate they will not notify their own sexual contacts;
- some states that have mandated confidential testing only (by name) are moving to an anonymous testing system.

The AIDS Policy Center is an AIDS legislative and policy research center that provides AIDS information and related research to legislators, policy-makers, and the media. In addition to publishing monographs on a variety of topics including state AIDS legislation, state AIDS funding, and general AIDS policy, the Center publishes a regular newsletter, *State AIDS Reports*.

For further information, contact Caitlin Ryan or Mona Row, AIDS Policy Center, Intergovernmental Health Policy Project, George Washington University, Suite 210, 2021 K Street, NW, Washington, DC 20006, tel: 206/676-8144.

Hypertension in Israeli Adolescents: Prevalence According to Weight, Sex and Parental Origin

MORDECHAI SHOHAT, MD, TAMAR SHOHAT, MD, MARC MIMOUNI, MD, MENACHEM NITZAN, MD, AND YEHUDA L. DANON, MD

Abstract: We studied the prevalence of hypertension in 57,499 male and 35,803 female Israeli military recruits and its relation with sex, weight, and parents' ethnic origin. The overall prevalence of systolic hypertension (>140 mmHg) was 1.75 per cent for males and 0.32 per cent for females. The prevalence of diastolic hypertension (>90 mmHg) was 0.41 per cent for males and 0.06 per cent for females. For males, the prevalence of systolic and diastolic hypertension increased with weight, exponentially. Males of Ashkenazi origin had a significantly higher prevalence of hypertension (systolic 2.52 per cent, diastolic 0.55 per cent) compared with those of Sephardi origin (systolic 1.12 per cent, diastolic 0.3 per cent). The prevalence of adolescents with systolic or diastolic blood pressure greater than the mean +2SD of each weight group ranged between 1.5–2.3 per cent. (*Am J Public Health* 1989; 79:582–585.)

Introduction

The prevalence of hypertension reported from different parts of the world ranges from 1 to 20 per cent of the adolescent population.[1] We report here the low prevalence among Israeli adolescents and the relation of hypertension to weight, parents' ethnic origin, and sex.

Methods

The National Service Act requires all Israeli boys and girls to present themselves at the age of 17 years at the induction centers for the purpose of classification for military service. Arab boys and girls are exempt from service and are not routinely examined in this system. Other exclusions are comprised mainly of girls who declare themselves to belong to orthodox religions (a proportion that dropped from 32 per cent in 1977 to 24 per cent in 1985)[2] and both boys and girls who are severely mentally retarded or handicapped. Ninety-four per cent of the boys born in 1966 underwent a complete physical examination at the age of 17 years and received a health classification score before the end of 1985.[3] Other characteristics of the population seen in the induction centers, the examinations, studies performed, and classification system are described in detail elsewhere.[3]

We analyzed the military recruit medical records of 58,499 healthy males (25,782 of Ashkenazi origin; 31,717 of non Ashkenazi [Sephardi] origin) and 35,803 healthy females (14,678 Ashkenazi; 21,125 Sephardi) who underwent a complete physical examination at the age of 17 years as part of a routine health assessment in the induction centers during a certain period of time between 1980 and 1986. Adolescents older than 17 years of age at the time of examination were not included in the study.

Blood pressure (BP) was taken once during the morning, using a mercury gravity sphygmomanometer (Baumanometer, W.A. Baum Company Inc., New York) with a standard adult cuff (bladder width 13 cm and length 24 cm) after the participant had been seated at least 5 minutes. All the measurements for blood pressure were taken by physicians. Weights were measured by the physicians in males and by trained soldiers in females. Weight was recorded to the nearest kilogram with the subject wearing only underwear. The physicians were instructed to determine the systolic blood pressure (SBP) and diastolic blood pressure (DBP) as the first and fifth phase korotkoff's sounds, respectively. During the next three to six months, adolescents with systolic blood pressure (SBP) exceeding 140 mmHg or diastolic blood pressure (DBP) 90 mmHg underwent five additional BP measurements on different days under simulated basal conditions (in the individual's family physician's office after the participant had been seated at least 5 minutes) and the mean of these measurements was recorded. Only those with a measurement mean pressure exceeding 140 mmHg systolic or 90 mmHg diastolic were considered hypertensive.[4]

After determining the prevalence of SBP and DBP for each sex according to parents' ethnic origin, we restricted our analysis to males weighing 46–90 kgs and females weighing 46–80 kgs. This selection included 95 per cent of the total recruit population of each sex and was aimed at preventing errors resulting from inappropriate cuff size. Data were classified by 5 kg subgroups and analyzed according to sex and parental ethnic origin.

Results

The crude prevalence of systolic hypertension (Table 1) was 1.75 per cent for males and 0.32 per cent for females, for diastolic hypertension, 0.41 per cent for males and 0.06 per cent for females. Males of Ashkenazi origin had a systolic prevalence of 2.52 per cent, diastolic 0.55 per cent, compared to a systolic prevalence of 1.12 per cent, diastolic 0.3 per cent

TABLE 1—Prevalence of Systolic (SBP > 140 mmHg) and Diastolic (DBP > 90 mmHg) Hypertension According to Sex and Ethnic Groups

	Ethnic Groups		
	% Ashkenazi	% Sephardi	% Total
Males			
SBP > 140 mmHg	2.52	1.12	1.75
DBP > 90 mmHg	0.55	0.30	0.41
Females			
SBP > 140 mmHg	0.35	0.32	0.32
DBP > 90 mmHg	0.11	0.09	0.10

From the Department of Pediatrics, Beilinson Medical Center, Petah Tiqva, Sackler School of Medicine, Tel Aviv University, and the Medical Corps, Israel Defense Forces. Address reprint requests to Dr. Mordechai Shohat, Department of Pediatrics/Genetics, Cedars-Sinai Medical Center, 8700 Beverly Boulevard, Los Angeles, CA 90048. This paper, submitted to the Journal January 20, 1988, was revised and accepted for publication August 9, 1988.

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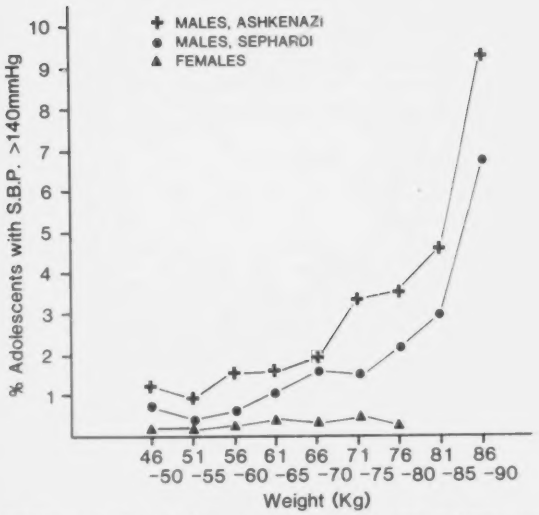

FIGURE 1—Per Cent of Adolescents (weighing 46–90 kg) Who Had SBP > 140 mmHg According to Weight Subgroups (values according to parental ethnic origin for the females were almost the same)

for those of Sephardi origin. The corresponding figure for females are Ashkenazi: systolic 0.35 per cent, diastolic 0.11 per cent; Sephardi: systolic 0.32 per cent, diastolic 0.09 per cent.

Figures 1 and 2 depict the per cent of adolescents (weighing 46–90 kg) who had SBP >140 mmHg and DBP >90 mmHg, respectively, according to weight subgroups and parental ethnic origin. Since the prevalence of both systolic and diastolic hypertension for females of both parental ethnic origins was similar in all weight groups, a single line depicts the mean values of the Ashkenazi and Sephardi female

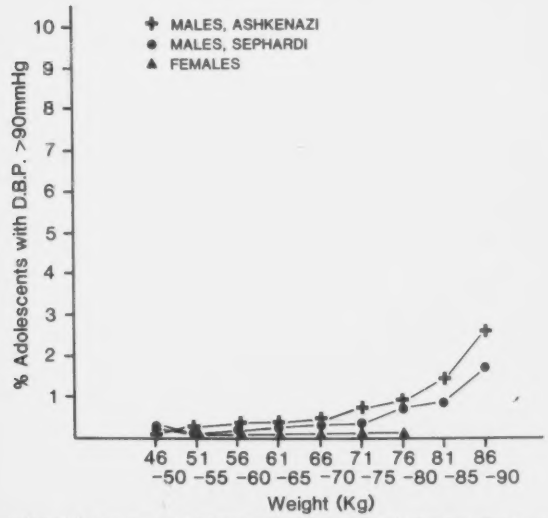

FIGURE 2—Per Cent of Adolescents (weighing 46–90 kg) Who Had DBP > 90 mmHg According to Weight Subgroups (values according to parental origin for the females were almost the same)

groups. For males overall, the prevalence of diastolic hypertension and systolic hypertension increased with weight exponentially. Controlling for comparable groups, systolic prevalences are higher in males than in females and higher in males of Ashkenazi origin than in those of Sephardi origin (Figure 1). Differences according to parental ethnic origin for the DBP (Figure 2) are less prominent.

Figures 3 and 4 depict the mean blood pressure values ±2SD for SBP and DBP, respectively, according to weight groups and parental ethnic origin. Diastolic blood pressures in females of both parental ethnic groups (Figure 3) were not significantly different in any of the weight groups and therefore Ashkenazi and Sephardi females are presented by a single (mean) line.

Discussion

Recently, a National Institutes of Health Task Force report presented guidelines for detecting high BP in children.[4,5] High BP or hypertension is defined as systolic and/or diastolic BP equal to or greater than the 95th percentile for age and sex on at least three occasions. For 17 year old male adolescents (WT: 84 kg), the cutoff points are 141 mmHg systolic and 91 mmHg diastolic, both of which are 1 mmHg higher than our definitions. The changes of BP

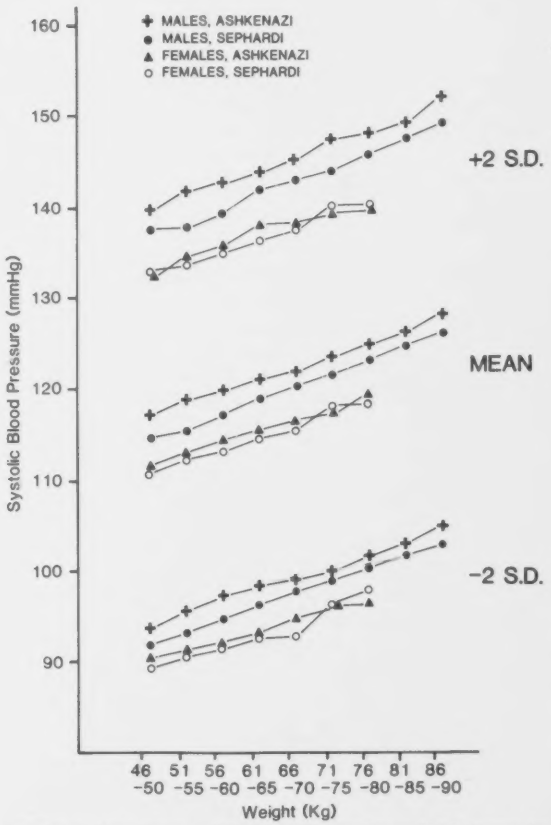

FIGURE 3—Mean Values + 2SD of the SBP According to Weight Groups for each Sex and Parental Origin

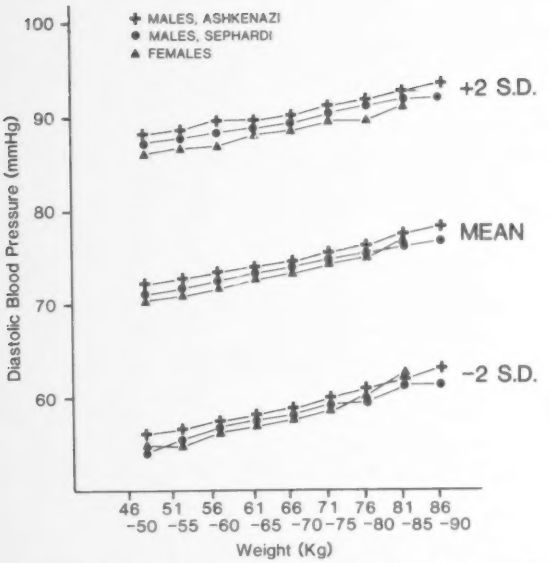

FIGURE 4—Mean Values +2SD of the DBP According to Weight Groups for each Sex and Parental Ethnic Origin (females of both parental ethnic origins are represented together as their values were almost identical)

according to weight in the adolescent group were not addressed in the report. The mean systolic and diastolic BP for males and females in our study was 3–5 mmHg higher than the new BP standards for adolescents of comparable weight. Our findings compare favorably with levels reported from South Carolina.[6] Nevertheless, the prevalence of hypertension was relatively low.

The true prevalence of hypertension in children and adolescents is not yet known. In his summary of the prevalence of hypertension reported from 13 regions in the world, Loggie[1] found that it ranged from 1 to 20.5 per cent. In studies using a definition of hypertension comparable to ours (SBP >140 mmHg, DBP >90 mmHg), the prevalence ranged from 1.4 to 11 per cent. Most of the studies reported are cross-sectional and repeat screening of those identified initially as "suspect hypertensive" is only infrequently reported.[1] Kilcoyne, *et al*,[7] and Reichman, *et al*,[8] who reported re-screening in New York City, found that significantly fewer "hypertensives" remained "hypertensive" at a second examination. Therefore, the fact that hypertension was defined in our study as the mean of five recurrent measurements on different days, under basal conditions, may explain the low prevalence of hypertension found in our study. Since the basal pressure in adolescents and young adults is a better predictor of hypertension in adulthood than "causal" BP,[9,10] it seems that the prevalence found in our study is closer to the actual prevalence of hypertension in adolescence. Since the causes for hypertension were not explored, the figures found represent the prevalence of hypertension, suggesting that essential hypertension is even less prevalent in our studied population. In adolescents who were under treatment for hypertension at the time of initial examination, blood pressure was determined under treatment. Although this might underestimate the blood pressure readings in those individuals, it probably changed the findings very little, since only a small number of adolescents (<0.02 per cent) were under such a treatment. It is possible that difficulties in excluding people from military service because of blood pressure elevations influenced the observing physicians to report lower readings and, therefore, reduced the prevalence of hypertension in our study. Nevertheless, since that bias in blood pressure measurement should not differ between ethnic groups and weight groups, cross-sectional comparison of distribution and mean values probably does reflect underlying ethnic differences.[3]

The American Heart Association committee recommends standardization of the bladder size in the sphymomanometer cuff.[11] For adolescents with arm circumference of 23–32 cm (equivalent to 46–90 kg of weight in the study and in other investigations), the standard adult cuff (13×24 cm) caused no more than a mean error of 5 per cent in the indirect blood pressure measurement.[12] BP measurement in our study suggests that there is an exponential increment in the prevalence of hypertension with increased weight in males regardless of parents' ethnic origin. The prevalence of systolic and diastolic hypertension among males weighing 86–90 kg is about 8–11 times greater than among males weighing 51–55 kg. For females, this increment seems to be less prominent. A positive correlation between hypertension and weight during childhood has been reported by other investigators.[13,14] This finding focuses on a major problem concerning the definition of hypertension. By what criteria should hypertension be defined? Should the fixed level (as in our study) be used or is it necessary to use specific percentiles (>97 per cent) for each weight group? In our population, the latter criteria would designate 1.5–2.3 per cent of the children as "hypertensive" for every weight between 46–90 kg.

Males of Ashkenazi origin have higher BP than those of Sephardi origin in almost all weight groups. As for females, this difference was less impressive. Several studies in Israel have shown ethnic differences in coronary heart disease (higher in Ashkenazi) as well as in other genetic characteristics such as blood groups and protein-enzymatic markers; these were reviewed recently by Goldbourt and Neufeld.[15] According to information from the Israeli Bureau of Statistics, the socioeconomic status of the Ashkenazi population is higher than the Sephardi whose diet is rich in salt. Adolescents of Ashkenazi origin are taller and heavier in comparison to those of Sephardi origin[16]; however, the differences in BP remained even after controlling for weight. Therefore, genetic factors may better explain this finding. A higher BP was obtained in American Black children and adolescents in comparison to Whites.[7,17] In order to decide if the same definition should be used for all races, it seems important to determine whether the prognosis of essential hypertension is the same.

ACKNOWLEDGMENTS

We are grateful to Dr. Jerome Kark and Ron Kedem for their help.

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AMSA Launches National Program to Teach Adolescents How to Resist Drugs, Peer Pressure

In response to concern over increasing drug use in the teen years and even earlier, the American Medical Student Association (AMSA) announced in late March the nation's first program that uses medical students to teach adolescents to resist peer pressure for alcohol and drug use.

Funded by the Office for Substance Abuse Prevention, US Department of Health and Human Services, AMSA's Substance Abuse Prevention program has trained medical students to provide prevention services to approximately 8,000 adolescents in grades four through eight. Between 150 and 200 students from eight or nine medical schools across the country have assessed the educational and personal needs of the teens and preteens in their local communities in order to design and conduct effective programs.

AMSA's President, Dr. Cindy Osman, noted the high probability that this program will have a profound effect on adolescents across the country. "This idea can work," she said. "We know that children and teenagers are most receptive when they learn from role models whom they respect. Young physicians can be powerful role models and we think they can get to kids before they begin to experiment."

Upon completion of the project, the American Medical Student Association will create educational materials which can be used in the classroom and other community settings for substance abuse prevention education.

The eight participating medical schools thus far include:
Brown University Program in Medicine, Providence, RI
George Washington School of Medicine, Washington, DC
Hahnemann University School of Medicine, Philadelphia, PA
University of South Florida College of Medicine, Tampa, FL
Louisiana State University School of Medicine, New Orleans, LA
Wayne State University School of Medicine, Detroit, MI
Washington University School of Medicine, St. Louis, MO
College of Osteopathic Medicine of the Pacific, Pomona, CA

For further information on the various student coordinators and faculty liaisons to the program, contact: Sharon McHale, American Medical Student Association/Foundation, 1890 Preston White Drive, Reston, VA 22091-4325. Telephone (703) 620-6600.

Results of Routine Restaurant Inspections Can Predict Outbreaks of Foodborne Illness: The Seattle-King County Experience

KATHLEEN IRWIN, MD, JANE BALLARD, MS, JOHN GRENDON, DVM, MPH, AND JOHN KOBAYASHI, MD, MPH

Abstract: To analyze the association between the results of routine inspections and foodborne outbreaks in restaurants, we conducted a matched case-control study using available data from Seattle-King County, Washington. Case restaurants were facilities with a reported foodborne outbreak between January 1, 1986 and March 31, 1987 (N=28). Two control restaurants with no reported outbreaks during this period were matched to each case restaurant on county health district and date of routine inspection (N=56). Data from the routine inspection that preceded the outbreak (for case-restaurants) or the date-matched routine inspection (for control restaurants) were abstracted from computerized inspection records. Case restaurants had a significantly lower mean inspection score (83.8 on a 0 to 100 point scale) than control restaurants (90.9). Restaurants with poor inspection scores and violations of proper temperature controls of potentially hazardous foods were, respectively, five and ten times more likely to have outbreaks than restaurants with better results. Although this study demonstrates that Seattle-King County's routine inspection form can successfully identify restaurants at increased risk of foodborne outbreaks, it also illustrates that more emphasis on regulation and education is needed to prevent outbreaks in restaurants with poor inspection results. (*Am J Public Health* 1989; 79:586–596.)

Introduction

Routine inspection of restaurants to prevent foodborne disease is mandated by food sanitation codes throughout the United States[1] and is recommended by the Model Standards for Community Health Practice of the US Public Health Service.[2] Although common sense dictates that the results of routine inspections should predict outbreaks of foodborne illness, this relation has never been studied in food service facilities other than cruise ships.[3] We therefore conducted a case-control study to determine whether routine inspection results and other characteristics were associated with reported foodborne outbreaks in restaurants using available data from Seattle-King County, Washington.

Restaurant Permit Program

For several decades, the Seattle-King County Department of Public Health has issued annual permits for permanent restaurants. Since January 1, 1986, when 3,076 restaurants had such permits,* characteristics noted on the permit have been entered into a computerized permit file.[4]

Restaurant Inspection Program

Sanitarians in five health districts use a standard reporting form developed by the Seattle-King County Department of Public Health for all routine inspections.[5] Data from the current form have been entered into a computerized inspection file since the form was adopted on January 1, 1986. The form identifies 42 types of violations classified as "critical" or "noncritical." Critical violations are thought to have a direct impact on foodborne disease, e.g., the temperature of potentially hazardous foods, food handling practices, the health status of food handlers (Table 1). Noncritical items are thought to play a minor role in foodborne illness, e.g., the cleanliness of nonfood contact surfaces, walls, and ceilings. A critical violation incurs a debit of 4-5 points from a perfect inspection score of 100, whereas a noncritical violation incurs a debit of 1-2 points. Violations are also classified as to type (food, food protection, personnel). After tallying all debit points, sanitarians assign a final score from 0 to 100 and a result category, which is largely a function of the final score. A score of 86 to 100 indicates a "satisfactory" result. A score of 70 to 85 or a violation of any critical item indicates an "unsatisfactory" result, requiring timely correction of violations. A score of less than 70 points indicates a "suspend permit" result, warranting permit suspension and restaurant closure. These score cutoffs were based on a county-wide study that simultaneously scored restaurants with the old 298-point inspection form and the new 100-point form; restaurants closed on the basis of the old form typically received scores of less than 70 on the new form.*

Foodborne Illness Investigations

Epidemiologists in the Seattle-King County Department of Public Health receive about 700 complaints of suspected foodborne illness each year.[7] All suspected outbreaks of foodborne illness are investigated to determine the number of affected persons, the symptoms of illness, the suspected vehicle, the food source, and the preparation, storage, or handling of food. A reported outbreak of foodborne illness is defined as an incident in which two or more persons have the same disease, have similar symptoms, or excrete the same pathogen after eating a common food or beverage. Poisoning by botulism or by a toxic chemical requires only one ill individual.[8]

Methods

Restaurants

Case restaurants were permanent restaurants with an active food permit in March 1987 and with a reported foodborne outbreak between January 1, 1986 and March 31, 1987. Although 36 restaurants were associated with an outbreak during this period, only 28 were permanent facilities with an active permit and therefore eligible for further analyses. Control restaurants were permanent restaurants with an active food permit in March 1987 and with no reported foodborne outbreak between January 1, 1986 and March 31, 1987. Two controls were randomly selected from eligible controls that had been matched to each case on health

Reprints are not available from authors. Kathleen Irwin, MD, is with the Center for Health Promotion and Education, Centers for Disease Control, Atlanta, GA 30333; Dr. Ballard and Dr. Kobayashi are with the Epidemiology Section, Washington State Division of Health; Dr. Grendon is with the Communicable Disease/Epidemiology Section, Seattle-King County Department of Public Health. This paper, submmitted to the Journal May 2, 1988, was revised and accepted for publication September 12, 1988.

*Unpublished data, Food Protection Program, Seattle-King County Department of Public Health.

TABLE 1—Risk of Foodborne Outbreaks Associated with Routine Restaurant Inspection Results, by Individual Violation, Seattle-King County, Washington, January 1, 1986–March 31, 1987

Number	Violation Type†	Debit Point Value	Odds Ratio	(95% CI)
	Food Violations			
1.	Foods from approved source, sound condition, not adulterated: no spoilage, no home canned foods†	5	3.2	(0.3, 36.6)
2.	Original container, properly labeled	1	1.0	(0.1, 18.9)
	Any food violation		2.1	(0.3, 13.1)
	Food Protection Violations			
3.	Potentially hazardous foods at safe temperatures during storage, display, service, transport, hot and cold holding (45° or below or 140°F or above)†	5	10.1	(2.2, 45.7)
4.	Potentially hazardous foods properly cooked to 140°F, except pork to 150°F, poultry to 165°F, and rare roast beef to 130°F†	4		(1.2,)*
5.	Potentially hazardous foods properly cooled, 4" food depth cooled to 70°F within 2 hrs to 45°F within 4 hrs; salads made with prechilled ingredients†	5	1.4	(0.4, 5.2)
6.	Potentially hazardous foods properly reheated to 165°F†	4	1.0	(0.1, 11.0)
7.	Enough facilities to maintain proper hot and cold temperatures, properly designed, maintained, operated. Potentially hazardous foods kept under temperature control except during necessary preparation procedures†	4	8.6	(1.0, 74.9)
8.	Thermometers provided and conspicuous	1	2.0	(0.7, 5.4)
9.	Potentially hazardous foods properly thawed	1	6.0	(0.6, 57.7)
10.	General food protection during storage, preparation, display, transportation, service; no double stacking; sneeze guards	2	2.4	(0.8, 7.2)
11.	Foods protected from cross-contamination during preparation and refrigerated storage. Foods not re-served†	4	1.7	(0.4, 7.0)
12.	Handling of food (ice) minimized; proper use of utensils	2	3.9	(0.4, 39.5)
13.	In-use food (ice) dispensing utensils properly stored	1	2.1	(0.5, 9.5)
	Any food protection violation		15.8	(2.0, 124.1)
	Personnel Violations			
14.	Personnel with infections or illness restricted†	5		(0.1)*
15.	Hands washed and clean; wash hands after using the restroom; after coughing, sneezing, smoking, eating, drinking; between handling raw and cooked; or otherwise contaminating hands. Good hygienic practices†	5		(0.6)*
16.	Clean clothes, hair restraints	1	NA	
17.	Food and Beverage Workers' Permits current for all personnel	1	1.8	(0.3, 12.0)
	Any personnel violation		3.3	(0.6, 18.0)*
	Food Equipment, Utensils Violations			
18.	Food (ice) contact surfaces: designed, constructed, maintained, installed, located	2	1.5	(0.4, 5.0)
19.	Nonfood contact surfaces: designed, constructed, maintained, installed, located	1	0.2	(0.04, 1.1)
20.	Food contact surfaces of equipment and utensils clean	2	1.8	(0.6, 5.4)
21.	Nonfood contact surfaces of equipment and utensils clean	1	0.6	(0.2, 2.2)
22.	Proper storage and handling of clean, sanitized equipment and utensils	1	14.9	(2.6, 85.4)
23.	Single-service articles properly stored and dispensed. No reuse of single-service articles	1	1.7	(0.2, 12.2)
	Any food equipment or utensils violation		1.8	(0.6, 5.8)
	Cleaning, Washing, Sanitizing Violations			
24.	Dishwashing facilities designed, constructed, maintained, located, operated (accurate thermometers, chemical test kits provided)	2	4.2	(0.8, 21.9)
25.	Equipment and utensils preflushed, scraped, soaked. Wash and rinse water clean, proper temperature	1		(0.03,)*
26.	Sanitization rinse: clean, proper temperature, concentration, exposure time. Equipment, utensils sanitized†	4	1.9	(0.6, 6.5)
27.	Wiping cloths clean, use restricted, stored in sanitizer	1	0.7	(0.2, 2.3)
	Any "cleaning, washing, sanitizing" violation		1.2	(0.5, 3.1)
	Water			
28.	Approved water source, hot and cold, under pressure; safe†	5	NA	
	Sewage, Plumbing Violations			
29.	Sewage and waste water disposed sanitarily. No cross-connection, back siphonage, backflow†	5	2.0	(0.1, 32.0)
30.	Plumbing installed, maintained	1	2.0	(0.4, 9.9)
	Any sewage and plumbing violation		1.5	(0.3, 6.7)
	Toilet, Hand-Washing Facilities Violations			
31.	Number, convenient, accessible, designed, installed†	4	3.6	(0.7, 19.9)
32.	Toilet rooms enclosed, self-closing doors, fixtures in good repair, clean; hand cleanser, sanitary towels/hand-drying devices provided, proper waste receptables	2	1.5	(0.5, 4.1)
	Any toilet and hand-washing facilities violation		1.6	(0.6, 4.2)
	Garbage, Refuse Disposal Violations			
33.	Containers/receptacles covered, adequate number, insect and rodent proof, pick-up frequency, clean	1	0.7	(0.1, 6.4)
34.	Outside storage area enclosures properly constructed, clean	1	4.0	(0.4, 44.1)
	Any garbage and refuse disposal violation		1.0	(0.2, 5.5)
	Insect, Rodent, Animal Control Violations			
35.	Presence of insects/rodents. No birds, turtles, or other animals†	4	6.5	(0.8, 51.1)
36.	Outer openings protected from flying insects/rodent proof	1	3.9	(0.4, 39.5)
	Any insect, rodent, animal control violation		3.8	(0.9, 16.1)
	Floors, Walls, Ceilings Violations			
37.	Floors constructed, clean, good repair, covered	1	1.5	(0.5, 4.6)
38.	Walls, ceiling, attached equipment: constructed, good repair, clean, smooth	1	2.0	(0.6, 6.8)
	Any floors, walls, ceilings violation		1.1	(0.4, 3.2)

(continued)

TABLE 1—Continued

Number	Violation Type†	Debit Point Value	Odds Ratio	(95% CI)
Lighting, Ventilation				
39.	Lighting provided as required, fixtures shielded. Rooms and equipment vented as required	1	1.2	(0.3, 5.0)
Other Operations				
40.	Toxic items properly stored, labeled, used†	4	1.9	(0.5, 7.4)
41.	Premises maintained free of litter; unnecessary articles; living/sleeping quarters separate; authorized personnel; dressing rooms, lockers	1	0.8	(0.2, 3.0)
42.	Clean, soiled linen properly stored	1		(0.1,)*
Any "Critical" Violation†			6.3	(1.8, 22.5)

*Odds ratio was indeterminate in matched and unmatched analyses; only the lower 95% confidence interval could be calculated in the matched analysis.
† "Critical" violations are numbers 1, 3–7, 11, 14, 15, 26, 28, 29, 31, 35, 40
NA—No restaurants had this violation

district and routine inspection date (±30 days), yielding 56 control restaurants for analysis.

Data

Data on each outbreak were collected from the Seattle-King County investigation files, including number of affected persons; implicated agent, vehicle, and contributory cause; and laboratory test results. We analyzed two variables from the permit file: type of ownership (corporate versus non-corporate) and seating capacity. All case and control restaurants had known values for at least one of these variables and were included in these analyses. Using the inspection file, we compared the routine inspection that preceded the outbreak for case restaurants with the date-matched inspection for control restaurants. We analyzed overall score, result category, specific violations, classes of violations, and inspection duration. One case and one control restaurant were excluded from the analyses because of incomplete inspection data.

In June 1987, managers of the 28 case and 56 control restaurants were telephoned to collect additional risk factor data. After obtaining informed consent, a trained interviewer questioned the restaurant manager on duty at the time of the call about restaurant characteristics, food preparation practices, employee turnover, training in food sanitation, and attitudes on food poisoning. After exclusions for non-response and interview refusal, 25 case and 48 control restaurants remained for analysis.

Mean scores, t tests, and confidence intervals (CI) were calculated using methods described by Ury.[9] Odds ratios and 95% CI for matched case-control analyses were calculated using PECAN.[10,11] Although we planned to match two control restaurants per case restaurant, exclusions during analysis necessitated a variable matching ratio. When sparse data made odds ratios in the matched analysis indeterminate, but at least one case restaurant and one control restaurant were exposed to a given factor, we calculated unmatched odds ratios and 95% CI.[12] When this was not possible, we applied exact methods to the matched data to calculate the lower 95% CI.[13]

Results

As shown in Table 2, the foodborne outbreaks affected from one to six persons with a mean of 2.9 persons. The implicated pathogen was unknown for most outbreaks; poultry was the most commonly implicated vehicle. The implicated contributory cause for most outbreaks was improper temperature control of food during cooking, cooling, reheating, holding, or storage. The interval between index inspection and outbreak ranged from 2.0 to 14.1 months, with a mean interval of 3.7 months. This interval is less than the four-month inspection interval recommended for all restaurants by the Seattle-King County Public Health Department.[14]

Case restaurants had a significantly lower mean inspec-

TABLE 2—Characteristics of Reported Foodborne Outbreaks in 28 Seattle-King County, Washington Restaurants, January 1, 1986–March 31, 1987

Characteristics	N
Number Persons Ill	
1	1
2	15
3–4	7
5–6	5
Total	28
Implicated Agent	
Unknown	22
Clostridium perfringens	2
Salmonella heidelberg	1
Shigella flexneri	1
Copper	1
Alkaline cleaner	1
Total	28
Implicated Vehicle	
Poultry	8
Rice	5
Fish	5
Beef	4
Pork	3
Beans	3
Other	4
Unknown	4
Total	36*
Implicated Contributory Cause	
Improper temperature control of potentially hazardous foods	25
Unsafe food source	5
Improper storage of toxic chemical	1
Cross-contamination	1
Poor food handler hygiene	1
Total	33*
Restaurant Cuisine Type	
American	9
Chinese	7
Mexican	7
Seafood	2
French	1
Japanese	1
Moroccan	1
Total	28

*Totals do not add up to 28 because in some outbreaks more than one vehicle or contributory cause was implicated.

TABLE 3—Factors Associated with Foodborne Outbreaks in Restaurants, Seattle-King County, Washington, January 1, 1986–March 31, 1987

Factors	Factor Present Cases/Controls*	Factor Absent Cases/Controls*	Odds Ratio	(95% CI)
Any improper food protection practice (violations 3–13)	25/31	2/22	15.8	(2.0, 124.1)
Improper storage or handling of equipment and utensils (violation 22)	6/1	21/52	14.9	(2.6, 85.4)
Potentially hazardous foods at unsafe temperature (violation 3)	15/10	12/43	10.1	(2.2, 45.7)
Any "critical" violation	22/24	5/29	6.3	(1.8, 22.5)
Inspection lasting ≥37 minutes	13/22	4/27	5.6	(1.1, 26.9)
Score ≤86 points	13/10	14/43	5.4	(1.5, 19.8)
Corporate owner	21/28	5/17	5.3	(1.1, 24.4)
"Unsatisfactory" or "Suspend Permit" result	20/22	7/31	3.9	(1.4, 11.0)
Restaurant size ≥150 seats	12/11	16/45	3.4	(1.1, 9.9)
Potentially hazardous food not cooked to proper temperature (violation 4)	3/0	24/53	†	(1.2, —)†
American cuisine specialty	9/32	16/16	0.2	(0.1, 0.7)

*Presented for unmatched data only.
†Odds ratio was indeterminate in unmatched analysis; only the lower 95% confidence interval could be calculated in the matched analysis.

tion score (83.8) than control restaurants (90.9) (difference = 7.1, 95% CI = −13.3, −2.18). Restaurants with an overall score of less than 86 were about five times more likely to have an outbreak than those with better scores. Restaurants that received an inspection result of "unsatisfactory" or "suspend permit" were about three times more likely to have an outbreak than those with "satisfactory" results (Table 3). Several specific violations significantly increased the risk of an outbreak, including any improper food protection practice (violations 3–13), especially improper temperature control of potentially hazardous foods (violations 3 and 4), improper storage and handling of equipment (violation 22), and any "critical" violation. Several other individual violations related to improper food protection practices had odds ratios of 2.0 but the confidence intervals included 1.0 (Table 1). Restaurants with inspections lasting longer than 36 minutes, with corporate owners, or with 150 or more seats were also more likely to have outbreaks than restaurants without these characteristics (Table 3). Two factors identified through telephone interviews were positively associated with outbreaks: Chinese cuisine specialty (OR=5.0, 95% CI=1.0, 25.8), and any Asian cuisine specialty (OR=4.0, 95% CI=1.0, 15.6). One factor, American cuisine specialty, had a clearly negative association with outbreaks (OR=0.2, 95% 0.1, 0.7) (Table 3).

Discussion

This study demonstrates that restaurants with poor routine inspection results were at increased risk of foodborne outbreaks. Key risk factors included a low score (less than 86 points), an inspection result warranting follow-up inspection or permit suspension, and violations of recommended food protection measures.

Our study was based on the inspection form used in Seattle-King County which differs from forms used by other locales and the US Food and Drug Administration[15] by assigning greater weight to violations of proper temperature controls of potentially hazardous foods. Most outbreaks in this series were probably caused by improper heating, cooling, cooking, holding, or storage of food—a finding consistent with a nationwide series of outbreaks reported to the Centers for Disease Control.[1] Inspection forms that give less weight to these temperature control factors may be less predictive.

It is widely believed that inspection results vary according to the sanitarian who performs an inspection.[14] We were unable to directly evaluate the quality of inspection data used in this study because Seattle-King County does not routinely require more than one sanitarian to do the same inspection. A recent evaluation in one district of the county where several sanitarians inspected the same restaurant at the same time suggested that the sanitarians were fairly consistent in identifying violations of proper temperature controls and cross-contamination but less consistent on overall score or the combination of violations accounting for that score.*

The strong association between outbreaks and improper storage and handling of equipment (violation 22) is difficult to explain on a biologic basis, and may be spurious because it is based on extremely sparse data. Conceivably, certain equipment violations, such as improper storage or handling of meat slicers, could be hazardous. Alternatively, improper use of equipment may reflect food-handling techniques more directly related to foodborne illness, a consideration if sanitarians do not use standard techniques for identifying violations.

Inspections lasting 37 minutes or more may have been associated with foodborne outbreaks because more time is needed to identify and record multiple violations. Large restaurant size may represent a risk factor simply because such restaurants serve numerous patrons, thus increasing the likelihood of finding two ill persons needed to identify an outbreak. Alternatively, large restaurants may be more likely to have an outbreak because of poor control of food temper-

*Unpublished data, Food Protection Program, Seattle-King County Department of Public Health.

atures, greater food volume, more complex menus, or less closely supervised food handlers. The association of the outbreaks with corporate ownership largely reflects the association of outbreaks with restaurant size. Compared with small noncorporate restaurants, large noncorporate restaurants were more likely (OR=5.0, 95% CI 0.5, 47.1) than small corporate restaurants (OR=2.5, 95% CI 0.6, 10.4) to have an outbreak.

We cannot rule out chance as an explanation for the positive associations with Asian or Chinese restaurants because the lower confidence bounds were 1.0, but these positive associations are plausible because certain food preparation practices in Cantonese-style restaurants have been found to be hazardous.[16] On the other hand, reporting bias could also explain these associations if Seattle-King County residents were more likely to report foodborne illness after eating in Asian restaurants than in other restaurants. Similarly, the apparent "protective" effect of American cuisine specialty could reflect less hazardous food preparation and sanitation practices or less intensive reporting of foodborne illness from American-style restaurants.

Poor inspection results should trigger appropriate education and regulatory action, such as follow-up inspection or permit suspension, which in turn should prevent outbreaks. Because our study illustrates that restaurants with poor inspection results are *more* likely to have outbreaks, it appears that the resulting regulatory action and education were not sufficient to prevent these outbreaks or that the restaurants did not adopt the recommended improvements on a long-term basis. Although we did not directly address this issue in our study, it probably reflects several problems. Restaurants with suspended permits are typically closed for less than 24 hours, a reprimand which might have little impact on a restaurant's profits or reputation. In addition, some restaurant managers complained during the telephone interviews that the education offered by sanitarians at the time of routine inspection is cursory or inconsistent. Finally, sanitarians in Seattle-King County and other locales[17] often report that food sanitation procedures taught at routine inspection have been abandoned by the next inspection.

Although this study demonstrates that the Seattle-King County inspection form can successfully identify restaurants at increased risk of foodborne outbreaks, it also illustrates that more emphasis is needed on regulation and education to prevent outbreaks in restaurants with poor inspection results. Permit suspension and timely follow-up inspections are clearly warranted when low scores or critical violations are noted. Detailed education to food handlers and their supervisors on the risks associated with specific violations is also needed. The risk estimates in Table 1 provide a simple instruction tool: for example, a sanitarian who notes unsafe storage temperatures of potentially hazardous foods (violation 3) could explain that this increases the risk of an outbreak ten-fold. Food protection programs should also assure that sanitarians use appropriate inspection techniques and that food handlers are certified in proper food preparation techniques. Finally, investigation of all complaints of foodborne illness will also help to monitor restaurants that pose an unusual public health risk.

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Multiple Program Participation: Comparison of Nutrition and Food Assistance Program Benefits with Food Costs in Boston, Massachusetts

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Abstract: The values of government cash and food assistance benefits are compared to estimated food costs in Boston for households whose sole source of income is Aid to Families with Dependent Children (AFDC) and who have housing in the private sector. Methods developed by the General Accounting Office (GAO) in 1978 are replicated. GAO found that the value of combined benefits can exceed the cost of the United States Department of Agriculture's Thrifty Food Plan. Key assumptions underlying GAO's methodology are challenged, and an alternate method is applied. The new results contradict GAO's conclusions and suggest that Boston food costs exceed the combined value of benefits that AFDC households may receive, although participation in multiple food assistance programs is more beneficial than receipt of benefits from single programs. The authors conclude that food stamps and AFDC benefits indexed to actual costs of living are needed to meet the food needs of low-income families in Boston. (*Am J Public Health* 1989; 79:591–594.)

Introduction

A variety of nutrition and food assistance programs operate in the United States, most of which provide direct benefits to low-income households meeting various eligibility criteria. Frequently, individuals and households eligible for one food assistance program qualify for others as well. Many low-income households also receive Aid to Families with Dependent Children (AFDC), the country's main means-tested cash transfer program. Although food stamp allocations for households are set by the federal government, AFDC benefit levels are set independently by each state. Both are based upon family size, income, and resources.

The use of multiple programs appears to be widespread. The President's Task Force on Domestic Food Assistance estimated that 76 per cent of food stamp households with children ages 5–18 years received free or reduced-price lunches in 1979.[1] Likewise, United States Department of Agriculture (USDA) figures have shown that 62 per cent of children receiving free school lunches are from food stamp households and approximately 42 per cent of food stamp households in the country also receive AFDC.[2] In Massachusetts, about 80 per cent of AFDC recipients receive food stamps.[3]

Some federal analyses have suggested that participation in multiple cash and food assistance programs may provide duplicative benefits to low-income households in excess of their food costs.[1,2,4] For example, in 1978 the General Accounting Office (GAO) published *Federal Domestic Food Assistance Programs: A Time for Assessment and Change*,[4] a report which examined both program administration and effectiveness. In one component, GAO compared the market value of different combinations of cash and food assistance benefits to USDA's Thrifty Food Plan and concluded that ". . . multiple participation . . . has created a situation in which benefits often exceed amounts needed. . . . for thrifty food plan diets."[4] GAO estimated that AFDC households might receive food assistance benefits equal to 230 per cent of the cost of the Thrifty Food Plan.

Other studies have suggested instead that multiple benefits are not excessive.[5–7] For example, the National Evaluation of School Nutrition Programs determined that school meals did not duplicate food stamps in their effect on household resources.[5] MacDonald examined participation in multiple programs using 1979 census information and found that 75 per cent of food stamp recipients and 85 per cent of food stamp eligible nonparticipants remained below the poverty line when the total value of income, cash, and food stamp benefits was considered.[6] In Massachusetts in fiscal year 1987, maximum AFDC and food stamp benefits for a three-person household totaled 83 per cent of federal poverty guidelines.

In this study, we re-examine the issue of participation in multiple programs in metropolitan Boston in the mid-1980s, using analytical methods similar to those presented in the 1978 GAO report.[4] We focus on households in private housing whose sole source of income is AFDC, the group most likely to have "excessive" benefits, according to GAO. In spring 1987, an estimated 10,000 Boston households met this description.[3] Although the GAO procedure is replicated, major assumptions underlying its methodology are challenged, and an alternative method for assessing the adequacy of food benefits is offered.

Methods

GAO Analysis

GAO created seven hypothetical low-income households that varied in size, composition, and ages of household members and determined the cash and food assistance programs for which each household was potentially eligible. One combination of program benefits was analyzed for each type of household and the monthly value of these benefits totaled, based on the reimbursable cost of child nutrition programs, the face value of food stamps, and 30 per cent of the monthly AFDC grant. These totals were compared to the cost of the USDA Thrifty Food Plan for each household. All data were based on program information for Oakland, California in spring 1976. The procedure was replicated for a sample of 95 actual Oakland, California households. GAO described seven of these households in terms of their composition and program participation in its report. Thus, GAO

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provided detailed information on a total of 14 households (seven hypothetical and seven actual).

Massachusetts Sample

The sample consisted of four households chosen from GAO's sample of 14. We excluded 10 of them because they did not receive AFDC (elderly, single males, or intact families) or they participated in a mix of benefit programs identical to one of those assigned to a household in our sample.

The compositions of the GAO households we examined were:

- Three persons: Mother age 29 and daughters ages 10 and 8.
- Four persons: Two households of similar composition:
  1) Mother age 40, sons ages 15 and 4, and daughter age 8.
  2) Mother age 25, daughters ages 7 and 4, and son age 6.
- Five persons: Mother age 32 and daughters ages 12, 10, 4, and 3.

Food Program Benefits—The programs we examined were AFDC, Food Stamps, National School Lunch (free meals), School Breakfast (free meals), and WIC (Special Supplemental Food Program for Women, Infants and Children) for our sample families. For comparable families, GAO examined AFDC, Food Stamps, National School Lunch (free and reduced-price meals), School Breakfast (free and reduced-price meals), Special Milk, and WIC. Because AFDC households would automatically be eligible for free school meals, we did not include reduced-price meals in our analysis. We chose not to include the Special Milk Program because virtually no Boston public schools sponsored it during FY'87 (regulations were still in effect that prohibited sponsorship by public schools that also offered one or more school meals).

Table 1 shows the value of AFDC and selected food assistance benefits for Massachusetts in FY'87. As in the GAO report, value was defined as the cash equivalent (AFDC and food stamps), the actual cost of the benefit (free school breakfast and WIC) to the government, or the actual cost plus the value of donated commodities (free school lunch). AFDC benefits shown are for families in private housing with no other source of income, and designated food stamp allotments are based on the AFDC grants. To determine the potential maximum contribution that food programs can make to food costs, we based our calculations of monthly values on daily participation, as did GAO.

TABLE 1—Value of AFDC and Selected Food Assistance Benefits (Monthly), Spring 1987

Program	Value
AFDC*	$491 (3 persons)
	$571 (4 persons)
	$652 (5 persons)
Food Stamps**	$140 (3 persons)
	$173 (4 persons)
	$200 (5 persons)
Free School Lunch†	$30.75
Free School Breakfast††	$14.75
WIC‡	$32.20

*AFDC: Maximum benefits for families of different sizes are shown, and include an additional $15 per month benefit for private sector housing.
SOURCE: Massachusetts Department of Public Welfare.

**Food stamps: Allotments were calculated for households of different sizes receiving maximum AFDC payments. The standard deduction of $99 and the maximum shelter deduction of $149 were subtracted to obtain net income on which the food stamps allotment was based.
SOURCE: Massachusetts Department of Public Welfare.

†Free school lunch: The federal reimbursement rate of 142.5¢ per meal plus 11.25¢ in commodities was multiplied by 20 school days per month for an estimated value of $30.75 per month.
SOURCE: Massachusetts Department of Education.

††School breakfast: The average federal reimbursement rate of 73.75¢ per meal was multiplied by 20 school days per month for an estimated value of $14.75 per month (rounded to $14.80).
SOURCE: Massachusetts Department of Education.

‡WIC: This is equivalent to the average monthly cost per WIC participant.
SOURCE: Massachusetts Department of Public Health.

While GAO considered one program participation mix for each of its households, we assigned several to each household in our sample.

Food Costs—Food costs were estimated in two ways, referred to as the Thrifty Food Plan (TFP) used by GAO, and the Estimated Household Food Costs devised by us. We estimated food costs in Boston indirectly, using data from the Bureau of Labor Statistics and the USDA. The cost of food purchased for home consumption in the Boston area, last measured in October 1981, was updated for use in this study.[8] The "intermediate" budget which most closely approximated average food costs[9] was updated by 19.1 per cent, the per cent change between the fall of 1981 and May 1987 in the Consumer Price Index for food at home in the Boston Metropolitan area (average cost changes for a food basket of cereals, produce, meat, and dairy products at selected stores).

The estimated monthly cost of food at home was $498.70 for the Bureau of Labor Statistics' prototypical family (mother, age 35; father, age 38; and two children, one under age 6 and the other between ages 6 and 15). To estimate monthly food costs for households of different compositions, the cost of the Thrifty Food Plan for each member of the family (taken from USDA data) was summed and the per cent of the total cost attributable to each member of the family was calculated. Results as a per cent of total household monthly food costs were: father, 29.9 per cent; mother, 27.1 per cent; older child 25.9 per cent; and younger child, 17.1 per cent. Multiplying these percentages by the May 1987 food at home cost estimate ($498.70) gave per capita results. On the basis of age distributions consistent with Thrifty Food Plan terminology, these figures were: father, age 20–50, $149.60; mother, age 20–50, $135.10; child through age 5, $85.30, and child age 6 through 16, $129.20 (figures rounded to nearest $0.10).

Results

We compared cash and food assistance benefits to food costs replicating the GAO analysis. Maximum monthly food assistance benefits (Table 1) were combined with 30 per cent of AFDC and compared to Thrifty Food Plan Costs (Table 2). Using the alternative approach, we compared maximum monthly food assistance benefits, excluding any contribution from AFDC, to Estimated Household Food Costs. Results are shown in Table 3. The results of the two types of analysis are compared in Table 4.

By the original GAO estimates, benefits for these households were as high as 203 per cent of Thrifty Food Plan Costs (Table 4, column 1). Using GAO's methodology and Boston data for FY'87, we found the value of government benefits included in the study provided 134.7 per cent to 194.5 per cent of the cost of the Thrifty Food Plan (Table 4, column 2). However, using the alternative approach (excluding AFDC and comparing the values of food assistance benefits to

TABLE 2—Comparison of Value of Selected Cash and Food Assistance Benefit Mixes to USDA Thrifty Food Plan (TFP) Costs

Size of Household	Potential Benefit Mixes (N of Participants)*	Monetary Value of Program Mix per Month**	Cost of TFP†	Per Cent of TFP Met By Assistance Program Mix
3 persons	AFDC(3) + FS(3)	$287.30	$213.30	134.7
	AFDC(3) + FS(3) + SL(2)	348.80	213.30	163.5
	AFDC(3) + FS(2) + SL(2) + SB(2)	378.30	213.30	177.4
4 persons††	AFDC(4) + FS(4)	$344.30	$257.40	133.8
	AFDC(4) + FS(4) + SL(2)	405.80	257.40	157.7
	AFDC(4) + FS(4) + SL(2) + SB(2)	435.30	257.40	169.1
	AFDC(4) + FS(4) + SL(2) + SB(2) + WIC(1)	467.50	257.40	181.6
4 persons‡	AFDC(4) + FS(4)	$344.30	$240.40	143.2
	AFDC(4) + FS(4) + SL(2)	405.80	240.40	168.8
	AFDC(4) + FS(4) + SL(2) + SB(2)	435.30	240.40	181.1
	AFDC(4) + FS(4) + SL(2) + SB(2) + WIC(1)	467.50	240.40	194.5
5 persons	AFDC(5) + FS(5)	$395.60	$297.80	132.8
	AFDC(5) + FS(5) + SL(2)	457.10	$297.80	153.5
	AFDC(5) + FS(5) + SL(2) + SB(2)	486.60	$297.80	163.4
	AFDC(5) + FS(5) + SL(2) + SB(2) + WIC(1)	518.80	$297.80	174.2
	AFDC(5) + FS(5) + SL(2) + SB(2) + WIC(2)	551.00	$297.80	185.0

*Abbreviations are: FS, food stamps; SL, free school lunch; SB, free school breakfast; and WIC, Special Supplemental Food Program for Women, Infants, and Children.

**Monetary values of benefits were derived from the values in Table 2.

†Monthly household food costs were estimated by using monthly per capita costs derived under "Methodology." These were: Mother age 20–50, $135.10; child through age five, $85.30; and child six through fifteen, $129.20. Thrifty Food Plan costs are from USDA, February, 1987.

††Mother age 40, sons ages 15 and 4, and daughter age 8

‡Mother age 25, daughters ages 7 and 4, and son age 6.

estimated household food costs in Boston), the data suggest that food stamps, school meals, and WIC are able to meet only 33.8 per cent to 66.3 per cent of estimated food costs (Table 4, column 3).

Discussion

The main differences between GAO methods and the new method derive from two assumptions. First, the GAO assumed that AFDC households would use 30 per cent of their grant for food. However, the Massachusetts Department of Public Welfare has shown that 63 per cent of AFDC recipients spend more than 75 per cent of their grant on rent.[10] Moreover, although the actual rents paid by AFDC households are not available, costs for private, unsubsidized rental units in Boston suggest that AFDC grant levels would not allow 30 per cent to be available for food. A Spring 1987 industry survey found that average monthly rents within Boston were $715 for a one-bedroom apartment (neighborhood averages ranged from $272 to $893) and $944 for a two-bedroom apartment (neighborhood averages $529 to $1260).[11] There were no vacancies in many of the less expensive neighborhoods. In contrast to these rents, Massachusetts AFDC benefits for three-, four-, and five-person households in Spring 1987 were $491, $571, and $652 per month, respectively (Table 1). AFDC benefits in Massachusetts rose 50 per cent from 1970 to 1985, whereas median Boston rents increased 329 per cent in the same period.[10]

Second, the GAO assumed that Thrifty Food Plan costs as given were an appropriate standard for measuring the adequacy of benefits. However, actual Thrifty Food Plan costs in Boston are below estimated food costs for all

TABLE 3—Comparison of Value of Selected Food Assistance Benefit Mixes to Estimated Household Food Costs

Size of Household	Potential Benefit Mixes (N of Participants)*	Monetary Value of Program Mix per Month**	Estimated Household Food Costs per Month†	Per Cent of Food Costs Met By Assistance Program Mix
3 persons	FS(3)	$140.00	$413.20	33.8
	FS(3) + SL(2)	201.50	413.20	48.8
	FS(3) + SL(2) + SB(2)	231.00	413.20	55.9
4 persons‡‡	FS(4)	$173.00	$478.80	36.1
	FS(4) + SL(2)	234.50	478.80	49.0
	FS(4) + SL(2) + SB(2)	264.00	478.80	55.1
	FS(4) + SL(2) + SB(2) + WIC(1)	296.20	478.80	61.9
5 persons	FS(5)	$200.00	$535.90	37.3
	FS(5) + SL(2)	261.50	535.90	48.8
	FS(5) + SL(2) + SB(2)	291.00	535.90	54.3
	FS(5) + SL(2) + SB(2) + WIC(1)	323.20	535.90	60.3
	FS(5) + SL(2) + SB(2) + WIC(2)	355.40	535.90	66.3

*, **, †Refer to footnotes in Table 2.

‡‡The two four-person households in Table 2 (†††) were combined in Table 3 because food costs and program eligibility were the same.

TABLE 4—Comparison of Current Results to 1978 GAO Report*

Size of Household	Per cent of food costs potentially provided by benefits: GAO Column 1	MA Dept of Public Health: Column 2 GAO Method	MA Dept of Public Health: Column 3 Alternate Method
3 persons:			
Mother age 29			
Daughters ages 10, 8	171	134.7–177.4	33.8–55.9
4 persons:			
Mother age 40			
Sons ages 15, 4			
Daughter age 8	166	133.8–181.6	36.1–61.9
Mother age 32			
Daughters ages 7, 4			
Son age 6	203	143.2–194.5	36.1–61.9
5 persons:			
Mother age 32			
Daughters ages 10, 12, 4, 3	166	132.8–185.0	37.3–66.3

*Current study shows a range because several program mixes were analyzed for each household. GAO analyzed one mix per household.

households shown in Tables 2 and 3. This finding is in keeping with data from the 1977–78 Nationwide Food Consumption Survey which indicated that 69 per cent of low-income households surveyed spent 110 per cent or more of their potential food stamp allotments on food.[12] A recent study by the Project Bread Food Stamp Hotline also found that food costs in Boston greatly exceed estimated Thrifty Food Plan costs.*

Thus we believe that GAO may have overstated the ability of government cash and food assistance benefits to meet actual food costs. Even though it is more beneficial for a family to take advantage of multiple nutrition and food assistance programs, maximum benefits from all these programs may not enable the most needy families in the Boston area to purchase an adequate supply of food. We conclude that food stamps and welfare benefits that are indexed to actual costs of living are needed if low-income families in the Boston area are to meet their food costs.

*Project Bread Hunger Hotline: Access to food: food prices, Food Stamps, and food availability in Boston. Manuscript in preparation, August 1987.

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Community-Based Nursing Administration: New Graduate Program at Penn

The University of Pennsylvania School of Nursing is developing a master's level educational program to prepare nurse managers for positions in community-based health care agencies. The effort is funded by a $160,000 grant from the W. K. Kellogg Foundation of Battle Creek, Michigan, and is managed by Cynthia Scalzi, PhD, RN, director of Penn's Nursing Administration Program.

The program will be designed with the assistance of a task force comprised of faculty and community leaders who will provide a broad base of community support for the program and assist in the development of a needs assessment survey.

The School of Nursing recognizes that nurse administrators today need superior planning and administrative expertise to handle the legal, regulatory, marketing, quality assurance, and financial issues that are unique to community-based health care programs.

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Fatal Violence among Spouses in the United States, 1976–85

JAMES A. MERCY, PHD, AND LINDA E. SALTZMAN, PHD

Abstract: In this paper we examine patterns and trends in homicides between marriage partners in the United States for 1976 through 1985 using data from the Federal Bureau of Investigation's Supplemental Homicide Reports (FBI-SHR). We identified 16,595 spouse homicides accounting for 8.8 per cent of all homicides reported to the FBI-SHR during this 10-year period. The rate of spouse homicide for this 10-year period was 1.6 per 100,000 married persons. The risk of being killed by one's spouse was 1.3 times greater for wives than for husbands. Black husbands were at greater risk of spouse homicide victimization than Black wives or White spouses of either sex. The risk of victimization was greater for spouses in interracial than in intraracial marriages and increased as age differences between spouses increased. From 1976 through 1985, the risk of spouse homicide declined by more than 45.0 per cent for both Black husbands and wives but remained relatively stable for White husbands and wives. Demographic patterns in the risk of spouse homicide were similar to those reported for nonfatal spouse abuse suggesting that the causes of spouse homicide and nonfatal spouse abuse may be similar. (*Am J Public Health* 1989; 79:595–599.)

Introduction

Intentional physical violence between spouses is a long-standing social problem not seriously addressed by researchers or policy-makers in the United States until the 1970s. The pervasiveness of violent events in marital relationships has since been convincingly documented.[1–4] Relative to the prevalence of abusive behavior between spouses, however, little attention has been given to the physical injuries associated with such behaviors, and there has been no attempt to undertake a population-based study of spouse homicide. To help address this gap in our understanding of the public health implications of spouse abuse, we undertook an investigation of spouse homicides in the United States for the years 1976 through 1985 to determine the demographic patterns for the risk of being murdered by one's spouse, and document annual trends in the rate of spouse homicide.

Methods

We obtained homicide data from the Federal Bureau of Investigation (FBI) for the years 1976 through 1985. On a monthly basis, state and local law enforcement agencies voluntarily report homicides and other major crimes to the FBI through the Uniform Crime Reporting Program (UCR).[5] The term homicide refers to those offenses reported to the FBI as murder and non-negligent manslaughter, i.e., "the willful (non-negligent) killing of one human being by another."[7] The classification of this offense is based solely on police investigation as opposed to the determination of a court, medical examiner, coroner, jury, or other judicial body. This study focuses on homicides in which the relationship of the victim to the offender was determined through police investigation as that of a legal or common-law husband or wife. We further focused on those homicides that police investigations categorized as being criminal (i.e., unjustifiable) and having only one offender.

Supplemental Homicide Reports (SHR) accompany UCR (Uniform Crime Reporting) reports on murder and non-negligent manslaughter include information on the sex, age, and race of victims and offenders, the victim-to-offender relationship, the circumstances under which the homicide occurred, and the geographic location of the homicide. We analyzed data by race only for Blacks and Whites because the low frequency of spouse homicide among other races makes it difficult to identify meaningful patterns.

Population estimates used to compute homicide rates were obtained from published data prepared by the Bureau of the Census.[8–15] Estimates of the "now married" population include only persons ages 15 years and older who were married for the first time, remarried, separated, or in common-law marriages. Therefore, this analysis was limited to homicides in which the victim was age 15 or older. Age-adjusted spouse homicide rates were calculated by the direct method of standardization,[16] with the 1980 United States "now married" population as the standard.

Population estimates used to calculate homicide rates for interracial and intraracial marriages, and by spouse-age differences, were only available from the 1980 census.[17] For categories of these variables, homicide rates were calculated by dividing the average frequency of homicide for the three-year period 1979 through 1981 by 1980 population estimates.

Results

In the United States from 1976 through 1985, 16,595 homicides involved persons killed by their spouse in the context of a single-offender criminal homicide (Table 1). The 16,595 spouse homicides exclude 340 single-offender homicides that involved an offender and victim who were divorced from each other, 193 homicides with multiple offenders, and 113 homicides that were classified as justifiable. These 16,595 victims constituted 8.8 per cent of all reported homicide victims in the United States during this 10-year period. In 98.6 per cent of these incidents, there was only one victim.

Sex, Race, and Age Patterns

The spouse homicide rate for this 10-year period was 1.6 per 100,000 married persons, with wives being at 1.3 times the risk of husbands (Table 2). Blacks accounted for 45.4 per cent of all spouse homicide victims. The rate of spouse homicide among Blacks was 8.4 times higher than that for Whites. White wives were at almost twice the risk of being killed by a spouse as White husbands, whereas Black wives had victim rates that were moderately lower than those for Black husbands. Age-adjustment of these rates left them virtually unchanged.

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TABLE 1—Number and Per Cent of Husband and Wife Victims by Category of Homicide, United States, 1976–85

Category of Homicide	Husbands		Wives		Total	
	Number	Per Cent	Number	Per Cent	Number	Per Cent
Spouse Homicides						
Single Victim/Single Offender	7,105	43.4	9,262	56.6	16,367	100.0
Multiple Victim/Single Offender	10	4.4	218	95.6	228	100.0
Total	7,115	42.9	9,480	57.1	16,595	100.0
Ex-Spouse Homicides*						
Single Victim/Single Offender	131	40.1	195	59.8	326	100.0
Multiple Victim/Single Offender	0	0.0	14	100.0	14	100.0
Total	131	38.5	209	61.5	340	100.0
Other Spouse-Related Homicides*						
Multiple Offender†	147	76.2	46	23.8	193	100.0
Justifiable	107	94.7	6	5.3	113	100.0
Total	254	83.0	52	17.0	306	100.0

*These categories of homicide have been excluded from subsequent analyses in this study.

†Includes three spouse homicides in which there were multiple victims.

Age-specific rates for Black husbands and wives exhibited a strong inverse association with age (Figure 1). Victim rates for Black wives and husbands peaked at 14.0 in the 15- to 24-year age category and then declined with age. Among Whites, however, age-specific spouse homicide rates did not decline consistently with age. Overall, racial differences in spouse homicide rates declined with age.

Spouse Differences in Race and Age

Of spouse homicides occurring from 1979 to 1981, 96.3 per cent occurred in intraracial marriages. Spouse homicides in marriages where the husband was Black and the wife was White constituted 1.4 per cent of the total, and marriages where the husband was White and the wife was Black accounted for 0.5 per cent. Spouse homicide incidence rates were 7.7 times higher in interracial marriages relative to intraracial marriages. The patterns in this risk varied by race and sex. White husbands, White wives, and Black wives were at greater risk of being killed by a spouse in an interracial than an intraracial marriage, whereas Black husbands were at less risk (Table 3).

The risk of spouse homicide victimization for both husbands and wives increased as the age difference between husbands and wives increased. Among couples where the husband was two or more years younger than the wife, the spouse homicide rate was 5.0 per 100,000 compared with 2.6 among couples whose ages were within 1 year of each other and 3.6 among those couples where the wife was two or more years younger than the husband. This positive association between the risk of spouse homicide victimization and spouse age differences held for husbands and wives of both races

TABLE 2—Spouse-Victim Homicide Rates* by Race of Victim for Husbands and Wives, United States, 1976–85

Race	Husbands	Wives	Total
White	0.7	1.3	1.0
Black	9.7	7.1	8.4
Total	1.4	1.8	1.6

*Rates are per 100,000 married persons.

Note: Excluded are 26 homicides for which information on race of victim was missing and 270 homicides occurring among victims of races other than White or Black.

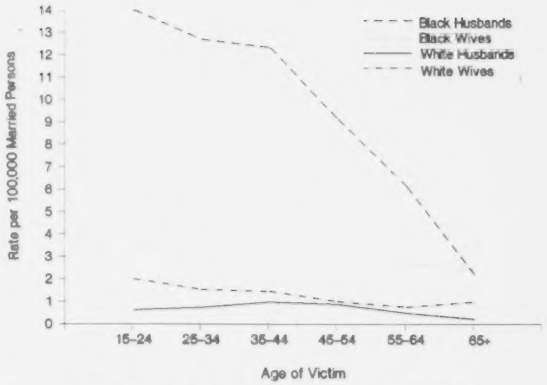

FIGURE 1—Spouse-Victim Homicide Rates by Race and Age for Husbands and Wives, United States, 1976–85

(Figure 2). This association also held regardless of the victim's or offender's age at the time of the homicide.

Weapon Involvement and Circumstantial Characteristics

Firearms were used in the perpetration of 71.5 per cent of spouse homicides from 1976 to 1986. Wife and husband victims were almost equally likely to have been killed by a firearm (71.0 per cent versus 72.2 per cent). Husband victims, however, were more likely to have been killed by a cutting instrument than wife victims (24.7 per cent versus 12.1 per cent), whereas wife victims were more likely to have been bludgeoned to death (11.8 per cent versus 1.6 per cent). White spouses were more likely to have been killed with a firearm than Black spouses. Among White victims, 83.2 per cent of the husbands and 72.5 per cent of the wives were killed with a firearm as compared to 63.9 per cent of Black husbands and 69.4 per cent of Black wives.

Arguments were noted as the circumstance most immediately associated with the killing in 67.2 per cent of spouse homicides. Arguments were more likely to be associated with the killing of husbands than wives (76.0 per cent versus 60.7 per cent) and the killing of Black spouses compared with White spouses (76.7 per cent versus 59.8 per cent).

TABLE 3—Rate† and Number* of Spouse-Victim Homicide by Racial Composition of Marriage for Husbands and Wives, United States, 1980

Race/Sex of Victim	Racial Composition of Marriage		Risk Ratio
	Interracial	Intraracial	
Husbands			
White	17.1 (14)	0.8 (1,002)	21.4
Black	8.9 (25)	12.2 (1,252)	0.7
Wives			
White	17.4 (49)	1.4 (1,784)	12.4
Black	13.4 (11)	9.6 (985)	1.4

*The number of homicides reported in parentheses under each homicide rate represents the frequency of homicides for a three-year period (1979–81). To calculate rates for this table the average annual frequency for the three-year period was divided by population estimates for 1980.

†Rates are per 100,000 married couples.

Note: Excluded are 10 homicides for which information on race of victim was missing and 74 homicides occurring among victims of races other than White or Black.

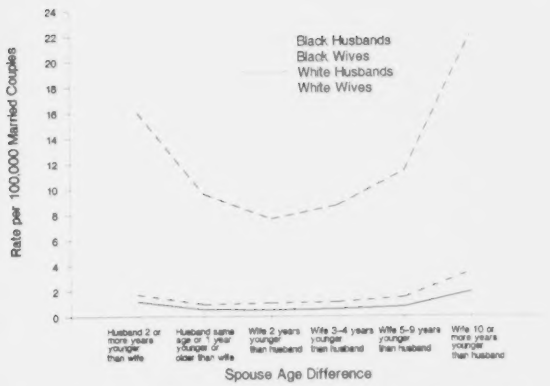
Black Husbands
Black Wives
White Husbands
White Wives
Rate per 100,000 Married Couples
Husband 2 or more years younger than wife
Husband same age or 1 year younger or older than wife
Wife 2 years younger than husband
Wife 3–4 years younger than husband
Wife 5–9 years younger than husband
Wife 10 or more years younger than husband
Spouse Age Difference

FIGURE 2—Spouse-Victim Homicide Rates by Spouse Age Differences and Race for Husbands and Wives, United States, 1980

Yearly Trends

Between 1976 and 1985, the spouse homicide rate dropped by 31.6 per cent from 1.9 per 100,000 married persons in 1976 to 1.3 in 1985. This decline in victimization rates was more dramatic for husbands than wives. The spouse homicide rate for husbands declined by 43.5 per cent; the rate for wives declined by 17.0 per cent.

When spouse homicide rates are further broken down by race, it becomes evident that the decline noted in the overall rate is almost entirely attributable to declines in the rates for Black husbands and wives. From 1976 to 1985, the spouse homicide rate for Black husbands declined by 52.0 per cent from 12.7 to 6.1, and the rate for Black wives declined by 45.8 per cent from 9.6 to 5.2 (Figure 3). In contrast, the rate for White husbands did not exhibit a consistent time trend, although from the beginning to the end of this 10-year period it declined from 0.7 to 0.5 per 100,000. The rate for White wives increased negligibly from 1.2 to 1.3 per 100,000. The risk of spouse homicide victimization for Black husbands relative to White husbands declined from 18.1 in 1976 to 12.2 in 1985. Similarly, the risk for Black wives relative to White wives declined from 8.0 in 1976 to 4.0 in 1985. Age-adjustment of these rates left the annual trends virtually unchanged.

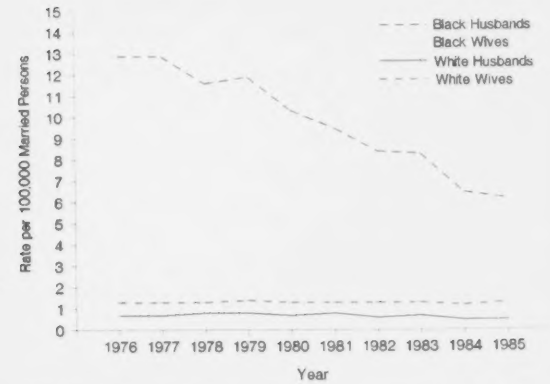

FIGURE 3—Spouse-Victim Homicide Rates by Race for Husbands and Wives, United States, 1976–85

Discussion

The finding that husbands and wives were nearly equal in the risk of spouse homicide victimization is also apparent in some studies of nonfatal spouse abuse. In both the 1976 and 1985 national family violence surveys, husbands and wives were at very similar risks of being physically abused by their spouses.[1,2] Other studies have reported similar findings.[4,18–21] Studies relying on police statistics, however, generally report that in most cases of nonfatal spouse abuse known to police the victims are women.[22–24]

Violent behavior by wives directed at their husbands may reflect acts of self-defense or retribution. Studies of homicides between intimates show that they are often preceded by a history of physical abuse directed at the women[25,26] and several studies have documented that a high proportion of women imprisoned for killing a husband had been physically abused by their spouses.[27,28] In a study comparing battered women who had killed their partners with abused women who had not killed their partners, Browne[29] found that the frequency and severity of wife abuse was associated with homicide. The weight of available evidence suggests that often wives kill their husbands in the context of a history of wife abuse.

Blacks had substantially higher spouse homicide rates and different patterns of risk than Whites. The risk of nonfatal husband and wife abuse has also been reported to be higher among Blacks than Whites.[1,30] Racial differences in the risk of spouse homicide, however, may only be a reflection of socioeconomic factors. Previous homicide research has demonstrated that socioeconomic factors are more important than race in explaining variations in homicide rates across aggregate units of analysis (e.g., states and cities).[31–33] Furthermore, other research indicates that variations in homicide rates among family and acquaintances may be more sensitive to socioeconomic factors than homicide rates among strangers.[34,35]

The strong inverse association found between age and the risk of spouse homicide for Black husbands and wives, but not for White husbands and wives, is a new finding. Research on nonfatal spouse abuse has also indicated an inverse association between age and risk,[1] but this finding has not been broken down by race. These patterns within Black marriages may reflect the well established association between violence and youth found in our society.[1] In addition, violent marriages may be more likely to dissolve through divorce, separation, or death than nonviolent marriages, thus reducing the chances that they would be as prevalent among older as younger couples.[1]

Differences between the race and age of spouses were found to be associated with a higher risk of homicide victimization. Although the risk of nonfatal spouse abuse has not been examined by differences in race or age, spouses in mixed religion marriages and wives in marriages where their educational or occupational level is high relative to their husband have been found to be at greater risk of physical abuse than spouses and wives in marriages without such differences.[1,20,36,37] It may be that persons with different racial or generational backgrounds bring very different outlooks and lifestyles to a marriage that may contribute to more conflict, greater stress, and, ultimately to violence. External pressures brought on by family ostracism or by people

outside the marriage, particularly for those partners in interracial marriages, could also contribute to stress and violence within the marriage.[1]

Firearms were used in the commission of most spouse homicides. In fact, firearms were more likely to be used in spouse homicides than nonspouse homicides (71.5 per cent versus 60.5 per cent). The presence of a firearm in the home may be a key contributor to the escalation of nonfatal spouse abuse to homicide.[38]

The risk of spouse homicide victimization declined between 1976 and 1985. This decline is almost entirely attributable to declines in the risk faced by Black spouses. Rates of severe nonfatal wife and husband abuse have been reported to have declined by 21.1 per cent and 4.3 per cent, respectively (Table 4).[2] These trends are in the same direction as those we found for spouse homicide. The decline in the risk of spouse abuse may be attributable to the spread of treatment programs and battered woman's shelters, the greater acceptability of divorce and changes in the structure of families (e.g., increases in age at first marriage), or to economic changes which may have decreased levels of stress on families.[2] These explanations, however, beg the question of why the decline in spouse homicide rates was limited to Black husbands and wives.

There are a number of actual and potential biases affecting both the number of spouse homicides and the estimates of the married population that should be considered when interpreting the results of this study. First, from 1976 through 1985 between 2 per cent and 4 per cent of the United States population was not served by law enforcement agencies participating in the FBI's Uniform Crime Reporting (UCR) Program.[5] In addition, approximately 4 per cent of the homicides reported to the FBI through the UCR were not accompanied by a Supplemental Homicide Report (SHR) form, which contains the information used in compiling the computer file used in this analysis. A second potential source of bias results from the FBI's monthly reporting requirement. For homicides occurring immediately before the FBI reporting deadline, the offender-to-victim relationship may be unknown at the time a report is made. It is possible, therefore, that some spouse homicides are hidden in the "undetermined relationship" category. A third potential source of bias is law enforcement use of inconsistent or vague criteria in identifying common-law marriages. There is no FBI reporting provision to apply a uniform definition for common-law status.

There are at least two potential biases in estimates of the population of married persons used in this study to calculate rates. Studies matching marital status as classified on death certificates with census records have found a tendency for widowed and divorced persons, especially between ages 15 and 44 years, to report their marital status to the census as married or single.[39–41] Meaningful differences in the self-reporting of martial status by race were not evident. Because of the large number of married persons between the ages of 15 and 44 years relative to the number of widowed and divorced persons, the effect of this bias on estimates of the population of married persons was judged to be negligible.[39] One other potential bias in estimates of the population of married persons concerns the ascertainment of common-law marriages in the census. Since the census relies on self-reports for information on marital status, respondents may define themselves as not married (i.e., single, divorced, or widowed), even though they may be in a relationship which is legally considered a common-law marriage.

Comparisons of data from studies of nonfatal spouse abuse to spouse homicide data presented here show a substantial degree of congruence between the demographic groups at high risk. Also, rates of both nonfatal and fatal spouse abuse show similar declines in the level of risk. These findings are consistent with the possibility that nonfatal spouse abuse and spouse homicide have a common etiology and that spouse abuse may escalate from less severe forms of physical abuse to forms that have a greater likelihood of causing injury or death. Thus, incidents of nonfatal spouse abuse and spouse homicide may be more appropriately viewed in the context of a series of violent events rather than as isolated events.[42] Repeated exposure to violence is likely to bring victims into repeated contact with a broad range of health and social services. Thus, contacts of abuse victims with emergency rooms, law enforcement agencies, social services, and physicians may provide a critical opportunity to prevent further injury or even death.

TABLE 4—Comparison of Changes in Spouse Homicide Rates* with Changes in an Index of Marital Violence between 1975 and 1985, United States

Violence Index	Year		Percent Difference
	1975/1976†	1985	
Spouse Homicide			
Wife Victim	2.00	1.70	−17.0
Husband Victim	1.70	.96	−43.5
Severe Violence			
Wife Victim	3800	3000	−21.1
Husband Victim	4600	4400	−4.3

*Rates for spouse homicide are expressed per 100,000 husbands or wives while rates for severe violence are expressed per 100,000 couples. SOURCE: Straus MA, Gelles RJ.[2] This index represent rates of nonfatal violent behaviors (i.e., kicking, biting, punching; hitting or attempting to hit with an object; beating; threatening with a knife or gun; or using a knife or gun).

†The spouse homicide rates reported in this table pertain to the year 1976, but the rates for the index of severe violence pertain to 1975.

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May Conference on "Meaning of the Holocaust for Bioethics"

The Center for Biomedical Ethics at the University of Minnesota will sponsor a conference on the "Meaning of the Holocaust for Bioethics," May 17–19, 1989, at the Radisson University Hotel in Minneapolis. Major themes include medical experimentation in the Third Reich, eugenics and euthanasia, the evolution of Nazi biomedical policies, the ethical status of Nazi research results, and analogies from the Nazi era in contemporary ethical debates.

Speakers include: Arthur L. Caplan (Director of the Center for Biomedical Ethics at the University of Minnesota), Jay Katz (Yale University), George Annas (Boston University), Robert Pozos (University of Washington), Benno Muller-Hill (University of Cologne), Peter Rossel (University of Copenhagen), Ivanyushkin Ye (Soviet Academy of Sciences), Ruth Macklin (Albert Einstein College of Medicine), and William Seidelman (McMaster University).

For further information, contact Continuing Medical Education, Box 202 UMHC, 420 Delaware Street SE, Minneapolis, MN 55455; telephone (612) 626-5525.

The Effect of Occupational, Marital and Parental Roles on Mortality: The Alameda County Study

PAMELA KOTLER, PHD, AND DEBORAH LEE WINGARD, PHD

Abstract: This study investigated the impact of combining marital, parental, and occupational roles upon 18-year risk of mortality from all causes. The respondents were 3,700 participants in the Human Population Laboratory cohort ages 35–64 who completed a comprehensive health and psychosocial questionnaire in 1965 and were followed for mortality status through 1982. Employment status and type of employment were not found to predict mortality risk among women. Contrary to the multiple roles hypothesis, there was virtually no impact upon mortality of increasing numbers of children among employed women, except possibly among single working parents. The major impact of children was felt by housewives who had significantly elevated risks when a child was present in the home or when they had four or more children. Neither the number of children nor the presence of a child in the home affected mortality risk of men. Controlling for a variety of factors thought to be related to mortality in a logistic regression analysis did not change the foregoing relation. (*Am J Public Health* 1989; 79:607–612.)

Introduction

Vital statistics show that women live longer than men.[1–4] One explanation proffered is that fewer women are employed outside the home, perhaps protecting them from risk factors of work. The last 30 years have seen a tremendous increase in women entering and remaining in the labor force: in 1950, 33 per cent of American women were employed; in 1985, the proportion had risen to 64 per cent.[5] The increase has been most dramatic among women with young children. Nevertheless, most cross-sectional[6–8] and prospective[9,*] studies show few differences in health between housewives and working women. Where differences exist, they tend to favor working women.[10–14]

Because most women continue to retain major responsibility for child care and homemaking when employed outside the home, it has been suggested that role overload and role conflict may adversely affect women's health.[2,15] Although employed mothers do experience increased stress and role conflict,[13,16,17] only one study has documented any prospective relation between multiple role responsibilities and morbidity.[18] In the Framingham Study, Haynes and Feinleib[18] found that while employed women and housewives did not differ in their risk of coronary heart disease, the risk of heart disease rose with the number of children among working women but not among housewives. Specifically, the excess risk of coronary heart disease was concentrated among ever-married clerical workers with children, particularly those married to blue-collar workers; this relation was not explained by standard coronary risk factors. Haynes and Feinleib theorized that the demands of paid employment and rearing a family may be excessive for some women. Verbrugge[19] also found that clerical workers in Detroit tended to have more symptoms and chronic health problems, and to use more prescription drugs than other groups of employed women.

An alternative view is that active involvement in the labor force may improve women's health, since employment can offer intellectual stimulation, productive activity, more social contacts and support, and greater income. These gratifications may outweigh stress arising from role overload or role conflict, particularly for women who are working by choice and not by necessity. Studies using data from the National Health Interview Survey tend to support this view; employed women with children were the healthiest group of women on a number of measures of morbidity, while women with few or no roles were among the least healthy.[19–21] These studies cannot discount the healthy worker effect—women who are in poor health tend to leave or never enter the labor force, leaving only the healthiest women in the labor force.

Although the issue of multiple role responsibilities has generally been treated as a women's issue, many men are involved in child care and household management and also have more responsibility as the size of the family increases. Another study of the Framingham data[22] found that men with three or more children had a risk of developing coronary heart disease 2.3 times that of men with no children. This effect remained after controlling for standard risk factors.

The current study examined the effect of multiple role responsibilities on the risk of mortality from all causes over an 18-year period among both men and women in Alameda County, California.

Methods

Study Population

In 1965, the Human Population Laboratory of the California State Department of Health Services surveyed a probability sample of Alameda County residents.[23,24] An extensive questionnaire, returned by 86 per cent of those contacted, included information on physical health status, health behaviors, employment status, family responsibilities, social contacts, and satisfaction with life. The present sample is restricted to the 1,731 men and 1,969 women who were between 35 and 64 years of age in 1965.

Mortality data were collected for the 18-year period 1965–82 for 97 per cent of the respondents.[23] A computer linkage was established with the California Death Registry and respondents were identified who died within the State of California. For those people who were not located and who had not died within the state, out-of-state death certificates

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*Waldron I: Relationships between labor force participation and women's health. Paper presented at the Annual Meeting of the American Psychological Association, Washington, DC, 1982.

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were sought from the state of birth. Through this process, 545 people were identified as having died during the interval 1965–82.

Women were categorized as working women (n = 726) if they had been employed outside the home fulltime for at least 50 per cent of their adult lives over the age of 22. Women were designated as housewives (n = 731) if they did not have the necessary years of work experience and if, in 1965, they indicated that they were housewives. The modal category of years worked at a fulltime job was 16–20 for the working women and never for the housewives. This strategy left 512 women who did not have the necessary years of work experience, but in 1965 were working fulltime or otherwise did not consider themselves to be housewives. These women were excluded from the analysis. All men ages 35–64 (n = 1,731) were included in the sample on the assumption that most men have been employed fulltime at least 50 per cent of their adult lives.

Variables

Type of occupation was coded into three major categories using the Dictionary of Occupational Titles: white collar workers (professionals, semi-professionals, technical workers, proprietors, managers, and officials); clerical and sales workers; and blue collar workers (foremen and craftsmen, operatives, apprentice craftsmen, service workers, laborers and private household help).[25] Type of occupation was not asked of women who said they were housewives in 1965, so only 567 of the working women responded to this question.

Marital status was categorized as single, married, or widowed/divorced/separated. Parenthood was assessed both as total number of children, and the presence or absence of a child between the ages of 3 and 19 in the home in 1965.

In addition to variables concerning role statuses, 12 variables considered risk factors in mortality studies were analyzed. These included four sociodemographic variables, two physical health variables, three health behaviors which have been established as independent risk factors, and three psychosocial variables. The categorization of these variables is presented in the Appendix.

Statistical Analysis

Age-adjusted relative risks of mortality and the Mantel-Haenszel chi-square statistic[26] were computed. The multiple logistic regression model[27] was used to estimate individual and joint effects of predictor variables on the risk of mortality simultaneously.

Results—

The sample studied consisted of 3,188 people ages 35–64 in 1965, 1,731 men and 1,457 women, of whom 726 were working women and 731 were housewives. Of the working women, 513 were employed in 1965 and 213 were not. As can be seen in Table 1, working women were far more likely than men to be clerical and sales workers and less likely to be white or blue collar workers. Fewer working women than men or housewives were married, had any children, or a child living in their home.

Age was the only sociodemographic characteristic on which working women and housewives did not differ. Working women had higher education, were more likely to be Black, and to have higher household incomes. When compared to men, working women were more likely to have completed high school. Men and working women had similar distributions of age, race, and income. Health behaviors, physical health status, and psychosocial characteristics were very similar for working women and housewives. However, working women were more likely to be current smokers and to drink alcohol (of any amount). Compared to working women or housewives, men were much more likely to drink alcohol, smoke cigarettes, and be physically active, and less likely to visit the physician and report poor physical health status.

TABLE 1—Prevalence* of Occupational, Marital, and Parental Roles among Men, Working Women, and Housewives in the Alameda County Cohort Ages 35–64 Years in 1965**

Role	Men		Working Women		Housewives	
	N	Per Cent	N	Per Cent	N	Per Cent
Occupation***						
White collar	584	33.8	152	26.8	...	...
Clerical, sales	202	11.7	240	42.3	...	...
Blue collar	938	54.4	175	30.9	...	...
Marital Status						
Single	72	4.2	61	8.4	4	0.5
Married	1527	88.2	474	65.3	653	89.3
Widowed, divorced, separated	132	7.6	191	26.3	74	10.1
Number of Children						
0	279	16.8	178	26.8	51	7.0
1–3	1084	65.3	406	61.1	488	67.1
4+	296	17.8	81	12.2	188	25.9
Child Present in Home						
Yes	963	58.0	306	46.0	449	61.8
No	696	42.0	359	54.0	278	38.2

*Prevalence = per cent distribution

**For all variables the distribution in men and housewives was significantly different from that in working women ($p < 0.001$) based on chi-square statistic.

***Working women (limited to 567 women answering this question).

Mortality varied by gender during the period from 1965 to 1982; 342 men and 203 women died. Age-adjusted 18-year risks of mortality from all causes showed that men had a higher risk of mortality than women; 19.8 per cent for men and 13.6 per cent for all women (RR = 1.5, CI = 1.2, 1.7). Among the women, housewives had a mortality risk of 14.2 per cent and all working women a risk of 13.5 per cent (RR = 1.05, CI = 0.8, 1.4). When housewives were compared to the 513 women who were currently employed in 1965, the relative risk was 1.2 (CI = 0.9, 1.7). As expected, mortality risks among both men and working women who were not employed in 1965 were substantially higher than the employed (RR = 2.4, CI = 1.9, 3.1 for men and RR = 1.7, CI = 1.1, 2.4 for women).

Figure 1 presents age-adjusted 18-year risks of mortality among men and women according to the type of work usually performed. Among working women, there were no substantial differences in mortality risk by type of work, nor did any of these groups differ from housewives. Among men, blue collar workers had the highest risk, but this was not substantially different from the risk of white collar workers (RR = 1.2, CI = 0.9, 1.4) or clerical and sales workers (RR = 1.2, CI = 0.8, 1.6). Among each of the three occupational groups, men had higher risks of mortality than did women, especially among blue collar workers (RR = 1.4, CI = 1.0, 2.0).

Adjustment for the effects of parenthood and the 12 variables shown in the Appendix had little impact on the pattern of mortality seen in Figure 1. The estimated relative risk for men compared to women was 1.7 (CI = 1.3, 2.1) after adjustment. The relative risk for working women compared to housewives was 1.04 (CI = 0.7, 1.5) after adjustment. We

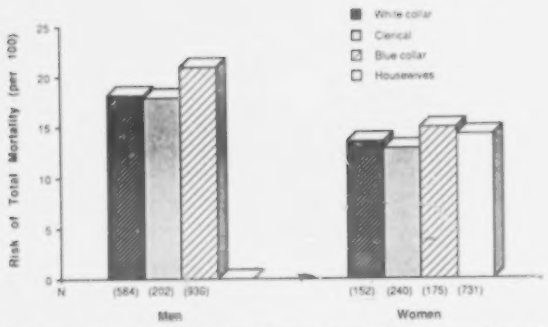

FIGURE 1—Age-Adjusted 18-Year Risk of All-Cause Mortality among Men and Women* by Type of Work
***with information available on type of work.**

examined the influence of each variable on mortality separately for men, housewives, and working women. Marital status was related to mortality risk in this study, especially for men (Table 2). Among all three groups, married people had the lowest risk; single people had intermediate risk, and widowed, divorced or separated people had the highest rates. (There were only four single housewives).

As seen in Table 2, number of children had no impact upon age-adjusted mortality risk for the men or working women. However, housewives with four or more children had a risk of mortality that was almost twice as high as that of housewives with fewer than four children (RR = 1.8, CI = 1.2, 2.5). Mortality risks did not differ by presence or absence of a child between the ages of 3 and 19 in the home for men, working women, or housewives.

Table 2 also presents relative risks for marital status combined with the presence or absence of a child in the home. A child in the home was associated with lowered risk among married working women, a raised risk among married housewives, and unchanged risk among married men. Among the widowed, divorced and separated, a child in the home was associated with a lower mortality risk for both men and housewives and a higher risk among working women. Small numbers of individuals in subgroups yielded unstable estimates of these latter effects.

Although neither occupation nor number of children showed an increased mortality risk among employed women, further analyses were undertaken to detect possible interactions, as was done in the Framingham study.[18] There were no differences between clerical and non-clerical workers for any of the marital/parental status groups. The comparison between ever-married clerical workers who had children and those who did not yielded a relative risk of 1.6 (CI = 0.7, 3.9). We found no increased risk of mortality among clerical workers with either three, four, or more children.

Since marital and parental statuses affected mortality risk differentially for men and women, we fitted separate logistic models for ever-married men and each group of women, including many lifestyle risk factors (Table 3). As seen in the unadjusted estimates, widowed, divorced, or separated men with no child at home had 2.7 times the mortality risk of married men with no child at home (CI = 1.7, 4.5). There were too few widowed, divorced or separated men with a child at home to get a stable estimate of risk.

Among the working women, those formerly married who had a child in the home were 2.2 times more likely to die than were formerly married working women who did not have a child in the home, but the estimate was unstable (95% CI = 0.9, 5.5). The presence of a child had little effect upon mortality among married working women.

The logistic model for housewives revealed that married housewives with a child in the home had 2.2 (CI = 1.1, 4.4) times the mortality risk of married housewives with no child at home. In addition, housewives with four plus children had 1.9 (CI = 1.1, 3.3) times the risk of housewives with fewer or no children, even after controlling for other risk factors.

Discussion

In this prospective study, long-term employment by itself was not associated with mortality. Employed women

TABLE 2—Eighteen-Year Age-Adjusted Relative Risks of All Cause Mortality among Men, Working Women, and Housewives by Marital Status and Parenthood: Alameda County, 1965–82

Characteristics	Men			Working Women			Housewives		
	Risk	RR**	CI***	Risk	RR	CI	Risk	RR	CI
Marital Status									
Married	18.4	1.0	. . .	12.2	1.0	. . .	13.4	1.0	. . .
Single	26.5	1.4	0.9–2.1	14.0	1.1	0.6–2.2	. . .	. . .	. . .
Widowed/divorced/separated	33.9	1.8	1.4–2.4	16.3	1.3	0.9–2.0	20.3	1.5	0.9–2.3
Number of Children									
0	20.9	1.0	. . .	12.0	1.0	. . .	11.9	1.0	. . .
1–3	19.0	0.9	0.7–1.1	14.0	1.2	0.7–1.9	12.2	1.1	0.6–2.3
4+	20.6	0.9	0.7–1.3	15.1	1.2	0.6–2.4	21.9	2.1	1.1–4.1
Child in Home									
No	21.1	1.0	. . .	14.1	1.0	. . .	14.4	1.0	. . .
Yes	18.5	0.9	0.7–1.1	14.2	1.0	0.6–1.4	18.9	1.4	0.9–2.2
Married in 1965									
No child	17.9	1.0	. . .	14.6	1.0	. . .	12.2	1.0	. . .
Child in home	18.7	1.0	0.8–1.3	7.6	0.6	0.3–1.1	18.7	1.6	1.0–2.8
Widowed/divorced/separated									
No child	35.8	1.0	. . .	13.7	1.0	. . .	40.2	1.0	. . .
Child in home	15.7	0.4	0.1–1.2	28.3	1.9	1.0–3.6	19.2	0.8	0.3–1.7
Number		1731			726			731	

*Risk = age-adjusted 18-year mortality risk per 100 people.
**RR = Mantel-Haenszel adjusted relative risk.
***CI = 95% confidence interval for relative risk.

TABLE 3—Results of Three Logistic Analyses for 18-Year All-Cause Mortality for Ever-married Men, Working Women, and Housewives in the Alameda County Cohort Ages 35–64 Years in 1965

	Men		Working Women		Housewives	
Variable (high/low risk category)	OR*	(CI**)	OR	(CI)	OR	(CI)
Age (per year)	1.1	(1.1–1.2)	1.1	(1.0–1.1)	1.1	(1.1–1.2)
Number of children (4+/0–3)	1.0	(0.7–1.5)	1.1	(0.5–2.4)	1.9	(1.1–3.3)
Marital status among non-parents*** (formerly married/married)	2.7	(1.7–4.5)	0.7	(0.4–1.4)	1.6	(0.6–4.2)
Marital status among parents*** (formerly married/married)	0.4	(0.3–20.7)	2.3	(0.9–5.4)	2.1	(0.2–1.3)
Child in home among married*** (yes/no)	1.2	(0.8–1.7)	0.7	(0.3–1.4)	2.2	(1.1–4.4)
Child in home among unmarried*** (yes/no)	0.2	(0.1–1.6)	2.2	(0.9–5.5)	0.6	(0.2–2.4)
Physical health						
(disabled/healthy)	2.9	(1.7–5.0)	1.3	(0.6–3.2)	2.7	(1.2–5.6)
(conditions/healthy)	1.3	(0.9–1.8)	0.7	(0.4–1.5)	1.0	(0.5–1.9)
(symptoms/healthy)	1.1	(0.7–1.6)	0.7	(0.3–1.4)	0.4	(0.2–1.0)
Smoking status (current/noncurrent)	2.1	(1.6–2.8)	1.5	(0.9–2.5)	1.6	(0.9–2.7)
Alcohol consumption (heavy/light)	1.4	(1.0–1.9)	2.7	(1.2–5.9)	0.9	(0.4–2.2)
(abstain/light)	1.1	(0.8–1.7)	1.8	(1.0–3.3)	0.9	(0.5–1.9)
Life satisfaction (dissatisfied/satisfied)	1.1	(0.8–1.4)	1.7	(1.0–2.9)	1.0	(0.6–1.6)
Contacts with friends (few/many)	0.9	(0.6–1.2)	1.8	(1.1–3.1)	1.6	(1.0–2.7)
Number		1659		665		731

No significant ($p < 0.05$) independent increase in risk was found due to race, education, income, occupation (included for men only), physical activity, visits to a physician or financial strain.

*OR = odds ratio

**CI = confidence interval

***Model included interaction term for marital status and presence of child in the home

and housewives did not differ in mortality risk and both groups had only 70 per cent the risk of men. This finding is in agreement with most cross-sectional studies.[6,7,*,11,12]

Unlike the Framingham studies,[18,22] clerical workers were not at increased risk in this study. In addition, clerical workers who had ever-married and had children experienced no increased risk when compared to ever-married nonclerical workers with children.

The results of this study may disagree with those of Framingham, first because the outcome in this study was mortality from all causes, rather than the incidence of coronary heart disease. However, the leading cause of death for women is heart disease and 32 per cent of the deaths among women in this sample were due to heart disease. Second, Framingham women were more likely to have left school after only eight years of education (20 vs 13.5 per cent), more likely to be employed in blue collar jobs (43 vs 30 per cent), more likely to be single (18.6 vs 8.4 per cent), and less likely to be formerly married (19.4 vs. 26.3 per cent). These variables may influence who becomes and remains a clerical worker in some way that leads clerical workers to have higher relative risks than nonclerical workers in Framingham but not in Alameda County. Another difference is that the Framingham study categorized any woman as a housewife who did not have the necessary years of work experience so that many women with substantial years of work experience (but less than half of their adult lives) who were working at the time of the study were still categorized as housewives. In this study, housewives were women without the necessary years of work experience who identified themselves as housewives in 1965. This may explain why an effect of parenthood was found for housewives in Alameda County and not in Framingham.

The time period of this study (1965–82) was one of rapid change in women's roles, both at home and in the labor market. In order to minimize the problem of change in roles after measurement in 1965, three strategies were used. First, the analysis was restricted to people who were at least 35 in 1965. Twenty years ago most women had completed their childbearing by age 35, so little change would be expected in terms of having more children. Second, the measure of employment was a lifetime work history in which women had to have worked outside the home for at least one-half of the years since age 22. Third, we examined the results for older and younger women (35–49 and 50–65) separately, on the theory that changes in parenthood, marital status, and lifetime work would be more likely in the younger group. There were no differences between the two age groups.

This study indicates that the effects of marital status and parental status upon mortality in women cannot be considered independently, but the fear that working women with children will suffer such severe role overload and role conflict that their mortality risk will rise appears unfounded. Married working women with children had the lowest mortality risks of any role status group in this study. Neither the number of children nor having a child in the home showed any relation to mortality risk among working women except possibly in the case of single parents. Although we could speculate about the overwhelming demands of being a single parent, this is a finding of marginal significance. These women were single parents in 1965; since that time the divorce rate has risen dramatically, making the role of single parent much less deviant. This, plus the greater availability of support services, and other major social changes that have taken place since 1965, may mean that the demands of being a single employed parent have decreased somewhat, and reduced the association with mortality.

The most surprising findings in this study concerned

housewives, among whom having four or more children or a child present in the home appeared to increase risk. One possible explanation is that the relative risk of coronary heart disease for women whose first pregnancy occurred before age 24 is 1.8 compared to women who had never been pregnant or whose first pregnancy occurred later.[28] If housewives with more children had earlier first pregnancies our findings may reflect this association, but direct information concerning age at first pregnancy was unavailable.

The apparent negative impact of numerous children upon women's health, particularly that of housewives,[7,13,28–30] deserves systematic investigation. Merely looking at the number of children or whether a child is present in the home is not sufficient; rather the biological, social, and psychological concomitants of having children need to be investigated.

Since women with the largest number of roles—employed, married women with a child present in the home—had the lowest risk of mortality in this study and women with the fewest roles—unemployed, unmarried women with no child at home—had a greatly increased risk of mortality, this suggests that involvement in a multiplicity of roles at least is not harmful to women's mortality risk. On the other hand, mortality is an extreme outcome and multiple roles may have a negative impact upon morbidity. As roles or statuses become more specifically defined, it becomes harder to rank order them in terms of demands. Does the housewife with five children have more or fewer demands than the working woman with two children? Where do the formerly married fit in? To examine fully the effect of combining work, marriage, and parenthood, it is necessary to design studies which actually probe the demands, conflict, benefits, and gratifications of each role rather than attempting to infer these from occupancy of the roles and to account for the likelihood that the degree to which women have chosen their roles will influence the outcome.

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(Appendix on following page)

APPENDIX
Categorization of 12 Variables Influencing Mortality: Alameda County, 1965–82

Variables	Categories	Remarks
Sociodemographic Variables		
Age	Years	
Race	White, Black, Other	
Education	0–8, 9–12, 13+ years	
Income	<$6,000, $6,001–9,999, $10,000–14,999, $15,000+	Household income adjusted for household size
Physical Health Variables		
Heath Status	No health problems, symptoms, chronic conditions, disability	Based on extensive list of questions concerning disability of movement and presence of chronic conditions, symptoms, and perceived energy levels
Physician Visits	0–9, 10+	Based on number of visits made to doctor during past 12 months
Health Behaviors		
Smoking Status	Current, not current	
Physical Activity	Inactive, active	Based on frequency of participation in five leisure time activities, swimming or walking, exercise, active sports, gardening, fishing or hunting and weighted by presumed energy expenditure
Alcohol Consumption	Abstain, light, heavy (≥drinks/month)	Based on frequency (number of times per week) and usual amount consumed
Psychosocial Variables		
Life Satisfaction	Dissatisfied, satisfied	Derived from nine questions concerning role satisfactions and feelings of pleasure in life
Contacts with Friends and Relatives	1–2 (few), 3–5 (many)	Derived from three questions on number of close friends and relatives and number of times seen per month
Financial Strain	0–4 (low), 5–7 (high)	Based on three questions asking about perceived strain with regard to income

Institute of Medicine Issues Call for Nominations for Lienhard Award

The Institute of Medicine is accepting nominations for the 4th annual Gustav O. Lienhard Award. The award, a medal and $25,000, recognizes individuals for outstanding achievement in improving personal health care services in the United States. The emphasis of the award is on creative or pioneering efforts that have appreciably improved personal health services rather than on the science base of health care. To encourage consideration of the widest possible range of candidates, there are no eligibility limits with respect to education and profession of individuals that may be nominated. Any individual or group may submit a nomination.

Names of nominees should be accompanied by a detailed written description of their accomplishments meriting this award. Only written material will be considered. Nominations must be postmarked by June 16, 1989, and should be submitted to:

Ms. Kay C. Harris
The Lienhard Award Committee
Institute of Medicine
2101 Constitution Avenue, NW
Washington, DC 20418

Support for the award is provided by the Robert Wood Johnson Foundation. Mr. Lienhard was chair of the Foundation's board of trustees from 1971 to 1987, a period in which the Foundation moved to the forefront of American philanthropy in health care. Previous recipients include: Julius B. Richmond, MD, architect of federal programs for Head Start and Neighborhood Health Centers in the 1960s and 1970s; Ernest W. Saward, MD, a pioneer in the establishment of prepaid group health plans; and Marie-Louise Ansak, a pioneer in nonprofit comprehensive care and founder of On Lok Senior Health Services, a day care center for the frail elderly.

Olfactory Function in Chemical Workers Exposed to Acrylate and Methacrylate Vapors

BRIAN S. SCHWARTZ, MD, RICHARD L. DOTY, PHD, CARL MONROE, MD, MPH, RICHARD FRYE, BA, AND SUE BARKER, RN

Abstract: An investigation of the olfactory function of 731 workers at a chemical facility which manufactures acrylates and methacrylates was undertaken using a standardized quantitative test. In a cross-sectional analysis of the data, no associations of chemical exposure with olfactory test scores were observed. A nested case-control study designed to evaluate the cumulative effects of exposure on olfactory function, however, revealed elevated crude exposure odds ratios (95% confidence interval) of 2.0 (1.1, 3.8) for all workers and 6.0 (1.7, 21.5) for workers who never smoked cigarettes. Logistic regression analysis, adjusting for multiple confounders, revealed exposure odds ratios of 2.8 (1.1, 7.0) and 13.5 (2.1, 87.6) in these same groups, respectively, and a dose-response relationship between olfactory dysfunction and cumulative exposure scores—semi-quantitative indices of lifetime exposure to the acrylates. The data also revealed decreasing exposure odds ratios with increasing duration since last exposure to these chemicals, suggesting that the effects may be reversible. (*Am J Public Health* 1989; 79:613–618.)

Introduction

A large body of animal toxicologic data indicates that many airborne chemicals can selectively damage cells within the olfactory neuroepithelium and/bulb, often in a specific and dose-related manner.[1–9] Depending on the chemical agent and the concentrations and durations of exposure, these changes may be either reversible or irreversible and may occur after either acute or chronic exposure.[6,9] Miller and others, for example, have reported that rats and mice exposed to ethyl acrylate, acrylic acid, and methyl methacrylate (5–300 ppm) evidenced, within the olfactory epithelium, respiratory metaplasia, squamous metaplasia, loss of olfactory neurons, hyperplasia of submucosal glandular elements, inflammation, degeneration, and focal necrosis.[10–13] These changes were dose-related to the severity and distribution of the damage and generally spared non-olfactory epithelia. A lower level of exposure below which damage consistently did not occur was not identified.

Despite these findings, the olfactory function of industrial workers exposed to such airborne chemicals has not been assessed, largely because an easy-to-administer and valid quantitative test of olfactory function has not, until recently, been available.[9,14,15] There are many cogent reasons why smell function should be assessed in the industrial setting:

- Normal olfactory function is important for worksite safety in a number of workplaces, e.g., during accidental leakage (odor detection thresholds can be lower than the level at which poisoning or explosion occurs).[9,16,17]
- Early detection of job-related changes in olfaction may allow reversal of detrimental effects (e.g., by decreasing exposure to the implicated agents).[18–23]
- Olfactory testing may be a useful addition to the neurotoxicologic testing armamentarium.[24,25] Quantitative tests of sensory and motor function have been employed but many of these tests require skilled technicians and are expensive, time-consuming, and invasive. Thus, the testing of olfactory function, as used in this investigation, has potential value in evaluating neurotoxicologic effects of chemicals that have detrimental effects on olfactory neurons and peripheral nerves; this testing is easy to perform, reliable, rapid, quantitative, and comparatively inexpensive.[15,26,27] Normative data are now available from thousands of subjects.[26]

The present study was undertaken to test the hypothesis that airborne ethyl acrylate, acrylic acid, and methyl methacrylate (all chemicals implicated to cause olfactory epithelial destruction in animals) cause a decrement in olfactory function in chemical workers after acute or chronic exposure. Both cross-sectional and nested case-control studies were performed.

Methods

Study Population

The study population consisted of all employees at a large manufacturing facility of the Rohm and Haas Company. The plant employs 909 workers in a variety of job activities and manufactures a number of chemicals, including acrylic acid and a variety of acrylates and methacrylates (especially ethyl acrylate, butyl acrylate, and methyl methacrylate). The facility operates 24 hours a day on three shifts. Participation in the study was voluntary, but encouraged through plant publicity from December 1986 to February 1987. Testing was conducted during two weeks in February 1987.

Approximately 850 of the 909 plant employees were available for testing during the two-week study period. The 59 workers who were unavailable for testing included workers absent for long-term disability, illness, and vacation. Of these 850 persons, 731 were tested; 710 (83 per cent of

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Carl Monroe died in December 1988. The authors gratefully acknowledge his primary role in the initiation and facilitation of this study. Dr. Monroe was an outstanding colleague who will be sorely missed.

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available employees, 78 per cent of total employees) were full-time Rohm and Haas employees, and 21 were contractors or consultants who happened to be on location during the testing period and desired to participate. The participants included 618 males and 113 females, ranging in age from 17–69 years (mean 42.9 years, SD = 11.3 years). The test group included 596 Whites, 98 Blacks, and 37 persons of Asian or Hispanic ethnicity. The subjects were from all three work shifts: 526 day shift, 131 evening shift, and 74 night shift.

Data Collection

The testing consisted of administering the University of Pennsylvania Smell Identification Test,* a questionnaire contained within the UPSIT, and a brief questionnaire to obtain shift and job information. The UPSIT consists of four booklets, each containing 10 microencapsulated odorants positioned on brown strips on the bottom of each test page.[14,26] The test is self-administered by scratching a pencil tip over the odorant label strip on the bottom of the page to release the fragrance, smelling the label, and choosing the perceived odor from four multiple choice answers. The test is forced-choice; the subject is required to choose one of the four answers even if no smell is perceived. The test is highly reliable and correlates with more lengthy test procedures.[14,27] Development of the UPSIT, as well as its relation to age, gender, and other factors, is described elsewhere.[14,26–29]

The test questionnaire obtained information on these study variables: age, gender, ethnic group, smoking history and dose, educational level, history of medical problems (self-reported problems which required treatment), history of smell or taste dysfunction, current medications, and work shift information. Detailed current job information was provided by the plant personnel office for all workers, and detailed employment histories were provided for the 154 workers in the nested case-control study.

Study Design

A cross-sectional (prevalence) study of olfactory function in chemical workers was undertaken which considered only current jobs and exposures. Next, a nested case-control study was performed to assess the cumulative effects of the chemicals of interest (see next section) on olfactory function. To be eligible for the case-control study, all cases and controls had to be full-time employees of Rohm and Haas for at least the past six months. Cases (N = 77) were defined as subjects scoring at or below the tenth percentile (for their age) on the UPSIT in 10 five-year age strata: ages 21–25, 26–30, 31–35, 36–40, 41–45, 46–50, 51–55, 56–60, 61–65, and 66–70 years. These cases were matched 1:1 to a random sample of controls defined as subjects scoring at or above the 50th percentile, for their age, in these same age strata. Cases and controls were matched on gender, ethnic group, and age.

It should be noted that the atypical selection of controls in this study (restricting eligible controls to the top half of their distribution) was designed to increase our ability to detect an association between chemical exposure and olfactory dysfunction if one existed. The design would tend to exaggerate the association (compared to a design using all non-cases as eligible controls, as is typically done) only if it existed; if no association between exposure and olfactory function existed, there would not be an increased risk of detecting one (i.e., no increased risk of a type I error).

Definition of Exposure

In the cross-sectional study, workers were classified into one of four exposure categories based on current job only:

- no significant chemical exposures (N = 319),
- exposure to chemicals other than acrylic acid, methacrylic acid, acrylates, or methacrylates (N = 193),
- exposure to lower levels of acrylic acid, methacrylic acid, acrylates, or methacrylates (N = 164),
- exposure to higher levels of acrylic acid, methacrylic acid, acrylates, or methacrylates (N = 55).

The foundation for the classification of workers into the lower and higher acrylate/methacrylate categories relied on the distinction between an Operator I and an Operator II in this plant. Those in the Operator I category spend most of their time in a ventilated control room monitoring the chemical process with little time spent outdoors on the chemical columns supervising workers or taking samples. An Operator II employee works primarily outdoors on the chemical columns taking samples and performing other duties. Industrial hygiene data collected in the plant from 1978–85 confirmed that Operator II employees and pumpers had the highest chemical exposure among workers in the plant. Pumpers work mainly to pump materials into and out of containers for shipping or receiving; during the opening of valves and lines for this purpose, high local concentrations ensue. When the distinction was unclear between Operator I and II in certain areas of the plant, Job Exposure Profiles and the aforementioned industrial hygiene data were used to classify workers into the lower and higher exposure categories. Job Exposure Profiles, developed by plant industrial hygienists, detail the chemicals worked with and the concentrations of exposure for many jobs in the plant, and thus were useful in this semi-quantitative assignment of exposure status. Finally, plant and corporate industrial hygienists helped to classify jobs whose categorization was unclear after the above steps were taken. These jobs were primarily ones at other Rohm and Haas plants and were approximately evenly distributed among the four exposure categories.

In the case-control study, exposure was studied in three ways:

- Exposure expressed as "never" or "ever" employed in job category 3 or 4 above for at least six weeks;
- the total duration of employment at the plant analyzed as a continuous variable; and
- a cumulative exposure score for each worker, analyzed as a continuous variable. The cumulative score was calculated by classifying each job title the worker ever held into one of the four exposure categories outlined above, multiplying the duration of time employed in each job title by the score for that job title, and adding these for all job titles ever held at the plant. The first two categories—no exposure and other chemical exposure—were assigned a score of zero, the lower acrylate/methacrylate category a score of one, and the higher acrylate/methacrylate category a score of two.

Current ethyl acrylate and acrylic acid levels varied from 0.01 ppm to 56 ppm; most areas of the plant have levels (as eight hour time-weighted averages) well below the threshold limit values (TLV) recommended by the American Conference of Governmental Industrial Hygienists of 10 parts per million (ppm) for acrylic acid and 5 ppm for ethyl acrylate,

*UPSIT; commercially available as the Smell Identification Test™, Sensonics, Inc, Haddonfield, NJ.

both as eight hour time-weighted averages. The highest levels were found when chemicals were pumped into containers for shipping, or when lines were opened for new connections or to obtain samples. Much of the available industrial hygiene data were for these specific high-exposure tasks (5–40 minute samples) and revealed levels at the high end of the aforementioned range. Current levels were used in determining exposure scores.

Analysis

Data analysis was performed on a microcomputer with the BMDP Statistical Software program. After detailed descriptive analyses, including frequencies and cross-tabulations, a three-way analysis of covariance with two covariates was used as the main model for the cross-sectional study.

In the case-control study, matched crude exposure odds ratios were calculated by discordant pairs,[30,31] and then conditional logistic regression was performed to calculate an adjusted exposure odds ratio. These two results were very similar to those calculated in unmatched analyses, so the matches were broken for the remainder of the data analysis. The only matched analysis results presented are the crude exposure odds ratio and a discordant pair table. In the unmatched analyses, crude odds ratios were calculated for all study variables, and then the Mantel-Haenszel procedure was used to calculate summary statistics after stratification by important confounders.[30,31] Logistic regression was then employed to adjust for other study variables in the calculation of the adjusted exposure odds ratio.

The nested case-control study examined three groups of workers: the entire group of cases and controls; only cases and controls who never smoked cigarettes; and cases redefined as those with UPSIT scores below 30 compared to all controls. The definition of cases below the tenth percentile by age allowed some cases with UPSIT scores as high as 35 out of 40; therefore, it was important to evaluate the worst-performing cases separately. A score of 30 or less on the UPSIT represents significant diminution of olfactory ability;[14] for example, based on normative data from thousands of subjects,[26] the approximate tenth percentile of performance on the UPSIT in the age range 16–50 years old is 34 for females and 33 for males.

Results

The cross-sectional analysis revealed mean (95% confidence interval) UPSIT scores of 37.8 (37.5, 38.0), 37.4 (37.0, 37.7), 37.0 (36.6, 37.5), and 37.6 (36.9, 38.1) in the none, other, lower acrylate/methacrylate, and higher acrylate/methacrylate exposure groups, respectively. The main model (three-way analysis of variance with two covariates) included ethnic group (White, Black, other), smoking status (never, previous, current), and exposure category (none, other chemicals, lower acrylates, higher acrylates) as grouping factors, and age and age-squared as covariates. This analysis revealed no differences in UPSIT scores in the four exposure groups ($F = 2.19$, $df = 3/698$, $p = 0.09$) after control of confounding by these variables.

Characteristics of cases and controls can be found in Table 1. There were apparent differences in the smoking history and ethnic categories. Results of the crude matched analysis are presented in Table 2. In the unmatched analysis, the case-control study revealed elevated crude exposure odds ratios for all workers and for never smokers (Table 3). A crude exposure odds ratio (95% CI) of 2.6 (1.2, 6.0) was found for cases with UPSIT scores below 30 (worst-performing cases (N = 34) compared to all controls, N = 111). Table 3 also shows the 2×2 contingency tables after stratification by smoking status. Effect modification appears to be present, with a markedly elevated exposure odds ratio in the never smoking group.

The logistic regression model used to adjust for confounding included all the measured independent variables and the matching variables: chemical exposure, smoking status (three categories), ethnic group (three categories), medications, age, history of smell dysfunction, history of medical problems, level of education, gender, and work shift tested. This model was then applied to the three groups of workers outlined above. There were no significant interactions with the exposure variable in this model for any of the three groups of workers. Control of all confounders with logistic regression revealed adjusted odds ratios of 2.8 (1.1,

TABLE 1—Characteristics of Cases and Controls

Characteristics	Cases (%)	Controls (%)
Number	77	77
Age (mean, SD)	42.8, 1.3	43.0, 1.2
Gender		
Male	64 (83)	64 (83)
Level of Education (mean, SD) (high school graduate = 12, college graduate = 16)	13.2, 0.3	14.0, 0.3
UPSIT score (mean, SD)	28.7, 0.8	38.9, 0.1
Medical History		
On medications	19 (25)	17 (22)
Medical problems	16 (21)	13 (17)
Smell problems	12 (16)	5 (6)
Ethnic Group		
White	52 (68)	58 (76)
Black	20 (26)	14 (18)
Other	5 (6)	5 (6)
Smoking History		
Never	29 (38)	30 (39)
Previous	16 (21)	28 (36)
Current	28 (36)	17 (22)
Unknown	4 (5)	2 (3)
Work Shift tested		
Day	54 (70)	61 (79)
Evening	13 (17)	13 (17)
Night	10 (13)	3 (4)
Exposure		
Ever exposed	47 (61)	34 (44)
Duration of employment, years (mean, SD)	15.5, 1.2	14.6, 1.2
Cumulative exposure score (mean, SD)	7.8, 1.2	6.6, 1.4

Percentages are shown in parentheses

TABLE 2—Crude Matched Exposure Odds Ratio

	Controls		
	Exposed	Unexposed	Totals
CASES			
Exposed	25	22	47
Unexposed	9	21	30
Totals	34	43	77

Crude matched exposure odds ratio = 2.4 (22/9)[30,31]
95% Confidence Interval = (1.2, 6.7)[31]

NOTES: Table contains 77 case-control pairs, comparing the exposure status of each pair. Odds ratio calculated by ratio of discordant pairs.

TABLE 3—Contingency Tables Stratified by Smoking Status

	Acrylate		
	Exposed	Unexposed	Totals*
A: Before Stratification			
Cases	47	30	77
Controls	34	43	77
Totals	81	73	154
Acrylate exposure odds ratio (95% CI): 2.0 (1.1, 3.8)			
B: After Stratification			
Never Smokers			
Cases	14	15	29
Controls	4	26	30
Totals	18	41	59
Acrylate exposure odds ratio (95% CI): 6.0 (1.7, 21.5)			
Previous Smokers			
Cases	10	6	16
Controls	18	10	28
Totals	28	16	44
Acrylate exposure odds ratio (95% CI): 0.9 (0.3, 3.2)			
Current Smokers			
Cases	23	5	28
Controls	11	6	17
Totals	34	11	45
Acrylate exposure odds ratio (95% CI): 2.5 (0.6, 10.0)			
Test for homogeneity for three stratum-specific odds ratios: p = 0.12.			

*Some subjects were missing smoking information.

7.0) for all workers (N = 141), 13.5 (2.1, 87.6) for never smokers (N = 57), and 2.9 (0.8, 11.3) for cases with UPSIT scores below 30 (N = 101). The most important confounders were smoking status, self-reported smell dysfunction, and educational level. The adjusted estimate of the overall exposure odds ratio was insensitive to changes in the method of controlling for smoking in the logistic regression model, e.g., when smoking was changed from a categorical variable to a continuous variable (in pack-years).

A summary of the data suggesting a dose-response relationship between olfactory dysfunction and the cumulative exposure score can be found in Table 4. Examination of these data reveals increasing exposure odds ratios as the cumulative exposure score increases, then a decrease in this odds ratio in the highest category of exposure.

Many employees worked in different areas of the plant where other chemicals were manufactured and acrylates and methacrylates were not used. In an attempt to control for the effects of other chemical exposures, a surrogate for exposure to these chemicals was calculated as years ever employed in other sections of the plant for all workers. The results of logistic regression analyses controlling for years employed in other areas of the plant, as well as all other measured independent variables, are presented in Table 5. There appears to be no confounding by these variables.

Finally, a summary of the data examining the reversibility of the olfactory dysfunction associated with these chemicals can be found in Table 6. The decreasing exposure odds ratios as duration since exposure increases suggests that the olfactory dysfunction associated with chemical exposure may be reversible. This finding is not entirely consistent, however, as can be seen by the adjusted odds ratios in Table 6. It should be noted that the categories in Table 6 were chosen for illustrative purposes; the results did not change when other cutpoints were chosen or when the year since last exposed variable was included in the models as a continuous variable.

TABLE 4—Evidence of a Dose-Response Relation with Exposure

Cumulative Exposure Score	No. Cases (%)	No. Controls (%)	Crude OR	Adjusted* OR
0	30 (41)	43 (59)	1.0	1.0
0.1–4.9	14 (54)	12 (46)	1.7	1.4
5–13	16 (64)	9 (36)	2.6	4.2
>13.1	17 (57)	13 (43)	1.9	2.4
		(C)oefficient	**SE‡**	**C/SE**
Logistic regression test for linear trend, adjusted for confounders:		0.3836	0.1863	2.059

*Controlling for all potential confounders with logistic regression (see text).
‡Standard error of beta coefficient

Discussion

This study evaluated exposure to acrylate and methacrylate vapors in prevalent cases of olfactory dysfunction and controls without olfactory dysfunction. There are many potential problems in such use of prevalence data in epidemiologic research.[31] Speaking in general terms, it is well known that numerous factors can influence the prevalence of disease (a generic term for the effect under study), such as disease duration and severity, that do not influence other effect measures such as risk or incidence. The association between exposure and disease can also be artifactual if disease duration is not identical in exposed and unexposed cases or if disease affects exposure status. The bias introduced by these factors can be in either direction (increasing or decreasing the real association) and is not predictable in this study. Review of employment records revealed evidence that persons employed in this plant enter one area (cyanhydrin, acetylene, acrylate, etc., processes with different physical areas in the plant) and stay in that area in various jobs for their entire career. This would argue against olfactory dysfunction being a prime determinant of exposure status.

The finding of the highest relative risk of olfactory dysfunction in the never smoking group is consistent with many plausible biological hypotheses. These subjects were

TABLE 5—Exposure Odds Ratios Controlling for Exposure to Other Chemicals

Controlling for Years Employed in:	Acrylate/Methacrylate Exposure Odds Ratio (95% CI)
Acetylene area	3.1 (1.2, 7.7)
Cyanhydrin area	2.8 (1.1, 7.0)
Laboratories	2.8 (1.1, 7.0)
Mechanical area	2.8 (1.1, 7.1)
Other areas classified as "other" chemical exposure, not listed above	2.9 (1.1, 7.6)
Total duration employed in mechanical area or any area classified as "other" chemical exposure	3.0 (1.1, 7.7)

TABLE 6—Evidence of Trend for Reversibility

Years since Last Exposure to Acrylates	No. Cases (%)	No. Controls (%)	Crude OR	Adjusted* OR
0–6.5	32 (60)	21 (40)	1.0	1.0
6.6–35	15 (54)	13 (46)	0.8	1.0
Never exposed	30 (41)	43 (59)	0.5	0.3
		(C)oefficient	**SE‡**	**C/SE**
Logistic regression test for linear trend, adjusted for confounders:		−0.5452	0.2485	−2.194

*Controlling for all potential confounders with logistic regression (see text).
‡Standard error of beat coefficient

not exposed to the insult of cigarette smoking on olfactory function, and thus would be expected to be at the highest risk for olfactory dysfunction from chemical exposure if this exposure caused olfactory dysfunction. This relative risk is substantial, with an odds ratio of 6.0 in the crude analysis and 13.5 in the adjusted analysis.

The cross-sectional study is best suited to evaluate the effects of recent exposure to these chemicals, and did not suggest any association between current job and olfactory dysfunction. The mean UPSIT score in the highest exposure group was virtually identical to the mean score in the non-exposed group. However, the nested case-control study, designed to evaluate the cumulative effects of lifetime exposure, revealed consistently elevated exposure odds ratios in many different groups even after control for confounding. The dichotomized exposure variable (ever or never exposed) and the cumulative exposure score can be criticized as being crude measures of exposure; however, these categorizations are likely to cause a nondifferential misclassification of exposure, which would tend to bias our findings toward the null.[31] Also, the assumption that specific jobs in the plant during the 1950s and 1960s had the same exposure levels as current jobs is unlikely to be true; again, this is a nondifferential misclassification of exposure. Finally, the definition of cases in this study as below the tenth percentile for their age in the cohort can be criticized as an arbitrary cut-off. This cut-off is also more likely to bias our results toward the null. In fact, when the analyses were repeated including just the worst-performing cases (those with UPSIT scores below 30), the crude exposure odds ratio was higher than for the overall group (the crude exposure odds ratio was 2.0 overall and 2.6 for the worst-performing cases); however, the adjusted estimate was unchanged.

The association between exposure and relative risk of disease is supported by the apparent finding of the dose-response nature of this relationship. The reduction in risk in the highest category of exposure remains unexplained. It may reflect a leveling-off rather than a reduction of risk, but further studies are necessary to clarify this finding. This dose-response relationship was not observed when duration of employment was substituted for the cumulative score in the logistic regression model. This important finding supports the contention that the cumulative score variable was not just a surrogate for duration of employment, but rather reflected a weighted score of exposure to the chemicals of interest. The finding of decreasing odds ratios with increasing duration since exposure is also consistent with many biological hypotheses; however, caution should be used in interpreting the meaning of this finding. Evidence obtained from prospective studies of olfactory function after exposure to these chemicals would be needed to conclude that the effects are reversible.

It should be noted that the plant manufactures a number of chemicals other than the acrylates/methacrylates to which workers are exposed. These other exposures were not measured but were controlled for in the analyses by controlling for years of employment in other areas of the plant. This method of controlling for other chemical exposures is acknowledged to be crude; it was the best method with the available data. The fact that the adjusted exposure odds ratio did not change after including these variables in the model reveals the absence of confounding by these "other exposures" variables. This was not surprising in view of the fact that workers in this plant tend to work in one area during their career; thus, work in other areas cannot be associated with exposure to acrylates/methacrylates, a necessary condition for confounding. It is possible, however, given the crude method of controlling for them, that other chemicals in the plant may be responsible for, or contribute to, the observed decrement in olfactory ability, alone or in combination with the acrylates/methacrylates.

As participation was voluntary, the 80 per cent participation could lead to another source of error in estimating the relation between exposure and olfactory dysfunction. The direction of this bias may be either to increase or decrease the association between exposure and olfactory dysfunction. This bias was evaluated by examining the jobs of the nonparticipants; with one exception, these jobs seemed to be a random distribution of nonparticipation by job title. The exception was that many workers employed in the mechanical group (electricians, pipefitters, carpenters, painters, etc.), approximately 40 per cent, chose not to participate. The mechanical group is a large, diverse group of workers, including those exposed and unexposed to acrylates/methacrylates. Although not obviously present, self-selection bias is a well-recognized occurrence in studies of this type.[31]

There is a large body of literature on olfactory dysfunction in humans secondary to chemical exposure; approximately 120 chemicals have been implicated to cause such dysfunction.[9,23] Most of the published studies are case reports or case series, few measure exposure quantitatively, and none employ rigorous epidemiologic methods. The studies that do measure exposure quantitatively suggest that hyposmia or anosmia only result from exposure to concentrations above the threshold limit value (TLV).[9,17] In the present study, the uncertainty in the exposure levels before monitoring began in 1978 causes difficulty in attributing the observed changes in olfactory function to a given exposure limit. This study cannot evaluate the effects of exposures orders of magnitude below the TLV (as may be encountered with environmental exposure). What is known about the physiology of the olfactory epithelium and its mucosal barrier would suggest that a threshold to the effects of chemicals may exist, in that this mucosal barrier absorbs and possibly degrades low levels of chemicals before they reach the olfactory neuroepithelium itself.

There are many points in the olfactory pathway that may be affected by exposure to these chemicals, and many pathophysiologic mechanisms may be postulated. The animal data revealed that exposure to acrylate and methacrylate vapors caused loss of olfactory neurons (neurotoxicity) as well as changes in the olfactory mucosa. The present study cannot determine the exact site of action of this exposure in humans; it may be caused by changes in the receptor site or

mucosa, neurotoxic effects, or central nervous system effects (olfactory bulb or pyriform, prepyriform, and entorhinal cortices). The animal data suggest that the most likely site is the olfactory mucosa or primary olfactory neuron.

Although this study suggested that exposure to acrylates and methacrylates is associated with a decrement in olfactory ability, as measured by a standard test of olfaction, it should be noted that this impairment is physiologic, but often not clinical. It is known that UPSIT scores in the range observed in the cases in this study are associated with a diminished ability to smell a number of environmental chemicals. The degree to which these changes effect the quality of life, lead to a decreased ability to perform job duties, or increase the risk of injury in areas where olfaction may be an early warning system for escape from hazardous airborne chemicals is unknown. The results of this work suggest that medical surveillance should be instituted to: clarify the effects of low-level exposure; determine the adequacy of the current exposure limits; determine whether the effects are reversible; and determine to what degree decreased olfactory ability affects quality of life and job safety.

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Hearing Loss among High School Farm Students

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Abstract: To study the prevalence of hearing loss among teen-aged farm children in central Wisconsin, audiometric threshold testing of 872 vocational agriculture students was carried out over a three-year period. The results indicate an increased prevalence of hearing loss among students actively involved in farm work, as compared to their peers not involved in farm work. Findings also suggest that use of hearing protection may reduce the risk of hearing loss among students who work on the farm, although few students report the use of such devices. These data also suggest that the hearing loss often observed in adult farmers may begin in childhood. (*Am J Public Health* 1989; 79:619–622.)

Introduction

Farmers are exposed to many sources of noise, often continuing for long periods of time at levels that are potentially damaging to hearing. A number of studies have suggested that farmers have higher than expected rates of hearing loss.[1–8] Many of these studies did not employ controls, while others have a built-in bias due to self-selection of subjects. Nevertheless, it is believed that the adult farm population suffers an increased risk of hearing loss, largely due to occupational noise exposure.

The present study was undertaken to determine whether there was greater evidence of early hearing loss among high school students engaged in farming than among students not involved in farming. A further goal was to describe sources of noise to which farm students are exposed, and to study the associations between various sources of noise, subject characteristics, and prevalence of early hearing loss.

Methods

From October 1985 through May 1988, audiologists from the Marshfield Clinic, Marshfield, Wisconsin, visited 12 area high schools to do audiometric threshold testing of vocational agriculture students. Students tested were between 12 and 19 years of age. Although all students were from vocational agriculture classes, they represented a wide range of participation in farming operations, from living and working more than 40 hours per week on the farm, to living in town and having no exposure to farm work. All students and their parents were asked to complete a brief questionnaire on health and hearing history, history of exposure to noisy farm and recreational equipment, and degree of participation in farm work. They were also asked to give their signed, informed consent to the testing procedures. At most schools, the entire class participated, except for the few students who were absent. To prevent temporary threshold shifts due to recent noise exposure, the audiologist described the project to the class and distributed ear plugs prior to testing. The students were instructed in the use of the ear plugs, and asked to use them when involved in any noisy activity either the night before or the day of testing. The vocational agriculture teachers encouraged complete participation, and later took responsibility for distribution of ear plugs to be used prior to testing. Each school was visited once per school year and many of the students were tested more than once during the 1985–88 period. For this report, only the first examination for each student was used, resulting in a total of 872 audiograms. Findings from subsequent audiograms collected in an ongoing longitudinal study will be described in a future report.

Testing was performed in a truck-mounted mobile testing unit with two specially designed Industrial Acoustics Corporation mini-testing booths. A Quest bioacoustic simulator, Model BA 201-25, was used for daily monitoring of background noise levels, and a General Radio octave band sound level meter, Model 1933, was used for assessing background noise levels. Two Maico manual audiometers, Model MA-39 with TDN-39 10 ohm earphones, were used for air conduction threshold testing. Regular electronic calibrations were conducted on each audiometer to ANSI-1969 standards.

A Hughson-Westlake modification of the ascending method threshold technique[9] was utilized for determining hearing levels. Hearing threshold was defined as where the tone was heard almost 100 per cent of the time and below which the tone was almost never heard. To minimize biases due to the learning effect, tones were presented in random fashion, both for ear tested and audiometric frequency (250, 500, 1000, 2000, 3000, 4000, 6000 and 8000 Hertz).

Analysis

For purposes of data tabulation, students were classified prior to testing into four distinct farming exposure groups based on their responses to the questionnaire and follow-up inquiries by the audiologist. These groups were:

- individuals who lived on farms and participated in virtually all farming activities (group A),
- individuals who did not live on farms, but worked on a farm either for hire or to help relatives or friends (group B),
- individuals who lived on farms, but did not participate or had minimal participation in chores and other farming operations (group C), and
- individuals who did not live on farms, and had little or no direct involvement with farm work (group D).

Hearing threshold values of 10 dB or less were classified as normal for each of the eight frequencies tested, and levels higher than 10 dB were classified as abnormal. This level was chosen based on a number of previous studies in similar age groups.[10–12] If a student had normal hearing (≤10 dB) at both 500 and 1000 Hertz, and abnormal findings (>10 dB) in the same ear at either 4000 or 6000 Hertz, he or she was

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considered to show evidence of early noise-induced hearing loss. A number of authors have suggested that this pattern of loss in the higher frequencies represents the early stages of noise-induced hearing loss.[4,8,13–15] This definition assigns subjects having below normal hearing at both low and high frequencies to the normal group, since this pattern is thought to be more characteristic of hearing loss secondary to illness or a middle ear disorder. Also, 26 subjects with mild hearing loss at low frequencies and more severe loss at high frequencies were classified as normal, and all but two of these possible misclassifications were in the groups with the most farm noise exposure (groups A and B).

Associations between demographic factors, sources of noise exposure, and prevalence of hearing loss were summarized using the odds ratio.[16] For multivariate analyses, logistic regression was used to estimate the odds ratio and its 95 per cent confidence intervals,[17] controlling for other relevant factors.

Results

Characteristics of the four farming exposure groups are shown in Table 1. The two groups with significant farm work exposure (labeled A and B) consisted primarily of males. The group reporting living on the farm, but having little involvement with farm work (labeled C) consisted primarily of females. Detailed probing by the audiologist of reported equipment use often revealed that some exposures were only occasional or secondary, as in the case of chain saws, where the student was frequently only a bystander, perhaps stacking or splitting wood. As another example, rifle use was often confined to the short deer hunting season and not a frequent activity. Although many of the responses to the questions on noise exposure shown in Table 1 are not thought to be reliable, collectively they suggest that the four exposure groups are similar with regard to health history, and that they differ markedly in the overall degree of exposure to noisy farm and recreational equipment.

Figure 1 displays the per cent of subjects in each of the four farming exposure groups who exhibit right ear hearing loss. At each frequency tested there was a higher rate of hearing loss in groups A and B, those with the most exposure to farm work. The differences (A/B vs C/D) are most marked at high frequencies. Similar results were found in both males and females, and also when hearing loss was defined as having threshold levels in excess of 20 dB (Figure 2). Hearing loss at the higher frequencies was observed even more often in the left ear.*

The percentages of students found to have evidence of hearing loss in at least one ear at either high or low frequencies were 71 per cent, 74 per cent, 36 per cent, and 46 per cent for groups A, B, C and D, respectively. Hearing loss in only low frequencies (500 or 1000 Hz) was uncommon, occurring in less than 5 per cent of the students. Hearing loss in only high frequencies (4000 or 6000 Hz) was much more common, occurring in nearly half of the students involved in farming (groups A and B) and in approximately one-fourth of the students not involved in farming (groups C and D). Hearing loss in both high and low frequencies was very infrequent in groups C and D, but was present in 13 per cent of right ears and 13 per cent of left ears in groups A and B combined.

Table 2 shows the prevalence of the combination of normal hearing in the low frequencies and hearing loss in the high frequencies, which we have labeled noise-induced hearing loss. The highest prevalence rates were found in the two groups most involved in farm work, with rates in groups A and B roughly twice those in groups C and D. The left ear was most frequently affected in group A, the students living and working on the farm, but there was little difference between left and right ear prevalence rates for the other groups. The prevalence of noise-induced hearing loss did not vary appreciably by month or by school.

Preliminary data from the ongoing longitudinal study indicate that the hearing deficits observed in the first year persist into the second year. On retesting one year later of students in groups A and B who were classified as having noise-induced hearing loss, 17 per cent had reverted to

*Detailed data available on request to author.

TABLE 1—Characteristics of the Study Subjects, by Farming Exposure Classification

Characteristics	Groups			
	A Live and Work on Farm	B Work on Farm, Live Elsewhere	C Live on Farm, Do Not Work	D Little or No Farm Exposure
Number of Subjects	445	198	50	177
Average Age (years)	15.8	16.0	16.3	16.1
Per cent Male	86	89	25	68
Per cent Reporting History of:	%	%	%	%
Ear Infection	37	38	33	36
Ear Surgery	4	9	0	1
Ear Drainage	14	15	14	16
Ear Pain	43	47	39	46
Ringing in Ears	43	52	31	47
Hearing Loss in Family	21	24	16	16
Per cent Reporting Use of:				
Tractor	94	78	61	24
Combine	35	16	12	2
Compressor	73	56	37	18
Barncleaner	84	52	51	13
Elevator	86	66	53	14
Haybine	72	52	35	8
Chopper	70	44	29	8
Silo Unloader	84	52	45	14
Grinder	62	44	29	12
Chain Saw	60	62	20	32
Motorcycle or Snowmobile	76	82	50	61
Pistol or Rifle	75	84	32	59
Woodworking Equipment	72	77	38	54
Amplified Music	79	82	82	81
Ear Protection	9	9	6	11

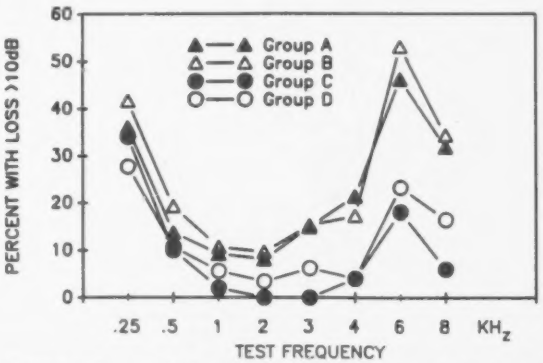

FIGURE 1—Per cent with Greater than 10 dB Hearing Loss at Each Frequency by Degree of Involvement in Farm Work, Right Ears

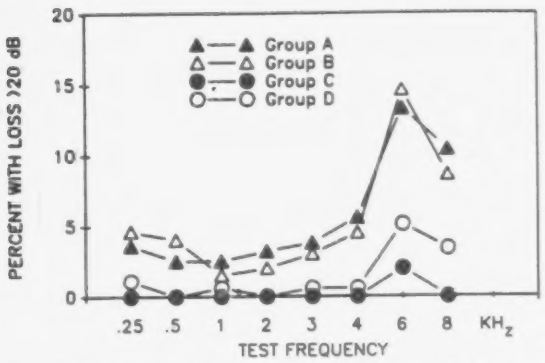

FIGURE 2—Per cent with Greater than 20 dB Hearing Loss at Each Frequency by Degree of Involvement in Farm Work, Right Ears

TABLE 2—Prevalence of Noise-Induced Hearing Loss* by Farming Exposure Classification and Ear Affected

Ear Affected by Noise-Induced Hearing Loss*	Farming Exposure Classification			
	A Live and Work on Farm n (%)	B Work on Farm, Live Elsewhere n (%)	C Live on Farm, Don't Work n (%)	D Little or No Farm Exposure n (%)
Neither	191 (42.9)	90 (45.5)	38 (76.0)	118 (66.7)
Left Ear Only	77 (17.3)	31 (15.7)	3 (6.0)	20 (11.3)
Right Ear Only	42 (9.4)	26 (13.1)	3 (6.0)	15 (8.5)
Both Ears	135 (30.3)	51 (25.8)	6 (12.0)	24 (13.6)
Either	254 (57.1)	108 (54.5)	12 (24.0)	59 (33.3)

*Defined as threshold at 500 and 1000 Hz. less than or equal to 10 dB and threshold at 4000 or 6000 Hz. greater than 10 dB in the same ear.

normal, versus 26 per cent of groups C and D. Of those previously classified as normal, 37 per cent of the students in groups A and B were classified as having hearing loss in the second year, versus 17 per cent in groups C and D. Overall, prevalence in groups A and B increased from 56 per cent to 63 per cent, and in groups C and D from 35 per cent to 37 per cent. When hearing threshold levels showed improvement, it was only rarely by more than five dB.

Table 3 displays results of a multivariate logistic regression analysis using additional independent variables to adjust

TABLE 3—Multiple Logistic Regression Analysis of Noise-Induced Hearing Loss in Either Ear, on Selected Subject Characteristics

Independent Variable	Estimated Odds Ratio	95% CI
*Group A (1 = Yes, 0 = No)	2.46	1.69, 3.59
*Group B (1 = Yes, 0 = No)	2.25	1.45, 3.49
*Group C (1 = Yes, 0 = No)	.61	.28, 1.30
Age (years)	1.05	.95, 1.16
Sex (1 = Male, 0 = Female)	1.02	.70, 1.50
Use of: (1 = Yes, 0 = No)		
Amplified Music	.81	.57, 1.15
Snowmobile or Motorcycle	1.24	.89, 1.74
Family History of Hearing Loss (1 = Yes, 0 = No)	1.24	.88, 1.76

*Group D = reference group.

TABLE 4—Multiple Logistic Regression Analysis of Noise-Induced Hearing Loss in Either Ear, on Selected Subject Characteristics for Students Involved in Farm Work (Groups A and B)

Independent Variables	Estimated Odds Ratio	95% CI
Group (1 = A, 0 = B)	1.22	.82, 1.82
Age (years)	1.04	.92, 1.18
Sex (1 = Male, 0 = Female)	1.02	.60, 1.74
Use of: (1 = Yes, 0 = No)		
Snowmobile or Motorcycle	1.39	.91, 2.14
Tractor	.96	.50, 1.85
Combine	.96	.63, 1.46
Compressor	1.18	.77, 1.81
Barncleaner	.58	.33, 1.02
Elevator	1.34	.78, 2.30
Haybine	.88	.52, 1.47
Chopper	1.13	.66, 1.94
Silo Unloader	1.10	.64, 1.90
Grinder	.74	.49, 1.12
Dryer	1.16	.74, 1.80
Hammering	.86	.58, 1.28
Family History of Hearing Loss (1 = Yes, 0 = No)	1.28	.85, 1.91
Use of Hearing Protection (1 = Yes, 0 = No)	.61	.34, 1.09

for the effects of age, sex, family history of hearing loss, and use of amplified music and snowmobiles or motorcycles. Memberships in groups A and B were most strongly associated with prevalence of noise-induced hearing loss. The estimated odds ratio of group A students to group D students was 2.5, as compared to an unadjusted odds ratio of 2.7. The estimated odds ratio of group B to group D students was 2.3, compared to the unadjusted ratio of 2.4. The analysis suggested that group C subjects were at a lower risk for noise-induced hearing loss than group D, but the confidence limits for this odds ratio include 1. Other variables included in this model appear to have only weak associations with prevalence of noise-induced hearing loss.

To further study the role of various sources of noise on prevalence of hearing loss, a separate logistic regression analysis is shown in Table 4 for students most heavily involved in farming operations (groups A and B). Students in group A had a slightly increased odds of hearing loss over students in group B, confirming the univariate results shown earlier (Table 2). Use of snowmobiles or motorcycles was associated with a slight but unstable increased odds of hearing loss, and the reverse was true with barncleaners. Use of hearing protection was suggestively associated with a reduced odds of noise-induced hearing loss, with an estimated odds ratio of .61 (95% CI = .34, 1.09).

Previous studies[1–4] have implicated tractor driving as a cause of hearing loss in farmers. Therefore another regression examined the subgroup of students in groups A and B who reported tractor use.* Most results parallel those shown earlier, but use of hearing protection in this group showed a more pronounced protective effect (O.R. = .53, 95% CI = .29, .96).

Discussion

Our data suggest that teenaged school children who are actively involved in farm work have increased prevalence of mild hearing loss and early noise-induced hearing loss (de-

*Data available on request to author.

fined as normal hearing at low frequencies combined with hearing loss in the high frequencies). Our results (coupled with reports of increased hearing loss in adult farmers) suggest that adult hearing loss may begin in childhood. Many farm children spend time in the vicinity of noisy machinery from an early age, including riding on and operating tractors and combines. In a small pilot study at our institution, we measured sound levels of tractors and other farm implements and found noise levels of many to be at or above levels for chronic exposure considered to be potentially damaging by current standards. Of 31 tractors tested at ear level in the driver's seat, without cabs or with a cab with an open window, only one tractor produced less than 85 dB at full throttle. Six of these tractors produced sound levels of 95 dB or more. As of this writing, OSHA (Occupational Safety and Health Administration) standards for maximum length of daily exposure to continuous or intermittent noise are eight hours at 85 dB and four hours at 95 dB. It is important to note that these standards are primarily meant to apply to adults. In addition to the noisy farm implements, people living in rural areas frequently make use of noisy recreational equipment such as firearms, power tools, snowmobiles, and motorcycles.

Our findings of greater left ear loss among farm students confirm those of other studies.[12,13,18,19] In the case of tractor drivers, it has been observed that the left ear is more likely to be exposed to engine noise, as drivers habitually observe their work by looking over their right shoulder, thereby shielding the right ear by the head shadow.[13]

The observed associations between the subject characteristics and prevalence of noise-induced hearing loss may have been affected by misclassification of farm students with significant noise-induced hearing loss into the normal group, due to the requirement that there be normal hearing at low frequencies. Also, the use of noisy equipment was recorded as a dichotomous variable, so that rare or occasional use of a piece of equipment was given equal weight to equipment used very frequently. Both of these sources of misclassification would be expected to weaken the observed associations. A third factor, for which we did not attempt to account, is the relation between the reported uses of the various farm implements. A student may typically be assigned a set of chores that could require the use of certain pieces of equipment and preclude the use of others. If there is much variation in the farm machinery used from student to student, and if most are significant sources of noise, the association of specific pieces of equipment with hearing loss would be difficult to untangle. It is also likely that even though students may not report operating certain pieces of equipment, they may have ridden on and been exposed to noise from these machines from an early age. The foregoing may explain the lack of any strong associations between reported use of farm equipment and prevalence of noise-induced hearing loss.

Aside from not participating in farm work, there are essentially two practical means presently available to reduce noise exposure on the farm. The first is use of hearing protection devices. Only 9 per cent of the farm students tested reported use of hearing protection when working in noisy areas; yet the evidence that students using hearing protection have a lower prevalence of hearing loss suggests there may be great potential for preventive efforts in this age group. A second means to reduce noise exposure is proper use of tractor and implement cabs. In our sample, only seven of the 563 tractor drivers exclusively used tractors with cabs. Our pilot study of noise levels found that tractors with cabs were less noisy, and that if the cab windows were closed, noise levels were reduced an average of seven dB more than if the windows were open. The present study showed only a weak protective effect of tractor cabs, however.

With the economic pressures facing many farmers, exposure of children to noisy and hazardous machinery could be expected to increase. Adolescent and teen-aged children often fill roles previously held by hired help. As farmers and their spouses take jobs off the farm to supplement income, farm children take on increasing responsibility for farm operations, and may be at increasing risk of hearing loss and other occupational health hazards of farming.

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Respiratory Disease Associated with Community Air Pollution and a Steel Mill, Utah Valley

C. Arden Pope III, PhD

Abstract: This study assessed the association between hospital admissions and fine particulate pollution (PM_{10}) in Utah Valley during the period April 1985–February 1988. This time period included the closure and reopening of the local steel mill, the primary source of PM_{10}. An association between elevated PM_{10} levels and hospital admissions for pneumonia, pleurisy, bronchitis, and asthma was observed. During months when 24-hour PM_{10} levels exceeded 150 $\mu g/m^3$, average admissions for children nearly tripled; in adults, the increase in admissions was 44 per cent. During months with mean PM_{10} levels greater than or equal to 50 $\mu g/m^3$ average admissions for children and adults increased by 89 and 47 per cent, respectively. During the winter months when the steel mill was open, PM_{10} levels were nearly double the levels experienced during the winter months when the mill was closed. This occurred even though relatively stagnant air was experienced during the winter the mill was closed. Children's admissions were two to three times higher during the winters when the mill was open compared to when it was closed. Regression analysis also revealed that PM_{10} levels were strongly correlated with hospital admissions. They were more strongly correlated with children's admissions than with adult admissions and were more strongly correlated with admissions for bronchitis and asthma than with admissions for pneumonia and pleurisy. (*Am J Public Health* 1989; 79:623–628.)

Introduction

On March 20, 1984, the US Environmental Protection Agency (EPA) proposed changes in the national ambient air quality standards for particulate pollution. Total suspended particulates (TSP) was to be replaced with a new indicator of particulate pollution that includes only those particulates with an aerodynamic diameter equal to or less than a nominal 10 micrometers (PM_{10}). On July 1, 1987, the EPA announced its final decision. The previous primary TSP standards were to be replaced, effective July 31, 1987, with a 24-hour PM_{10} standard of 150 micrograms per cubic meter ($\mu g/m^3$) with no more than one expected exceedance per year and an annual PM_{10} standard of an expected arithmetic mean of 50 $\mu g/m^3$.[1]

Earlier studies of the health effects of particulate pollution[2–8] revealed a possible connection between air pollution and human health, and launched a wave of research exploring this connection.[9–19] Recent research has observed that even moderately elevated concentrations of particulate pollution may result in reductions in children's pulmonary function[20,21] and increased risk for bronchitis and other respiratory illnesses.[22] Other recent research questions the existence of a threshold level.[1,23]

Previous studies have not used PM_{10} as an indicator of particulate pollution. Recent experiences in Utah County have provided a unique opportunity to investigate a possible association between respiratory health and different levels of PM_{10}. Utah Valley has had daily monitoring of PM_{10} since April 1985; it has an extremely low percentage of smokers; it has experienced a prolonged shut-down and then reopening of the steel mill, its largest source of particulate pollution; over time, since monitoring of PM_{10} began, the valley has experienced considerable variability in levels of fine particulate pollution; and hospital inpatient admissions data for respiratory illnesses can be obtained. The objective of this paper is to report what has been observed in Utah Valley with respect to hospital admissions for respiratory illnesses and PM_{10} levels.

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Methods

Study Area

Utah Valley, located in Utah County of Central Utah, is the third largest county in the state with a population of 258,000 in 1987.[24] Approximately two-thirds of the population resides in five nearly contiguous cities situated on a valley floor with an elevation of approximately 1,402 meters above sea level bordered east and west by mountains (Figure 1).

Based on an unpublished 1986 Utah State Department of Health survey, only 5.5 per cent of Utah County's adults (18 years of age or older) smoke; approximately 90 per cent of its

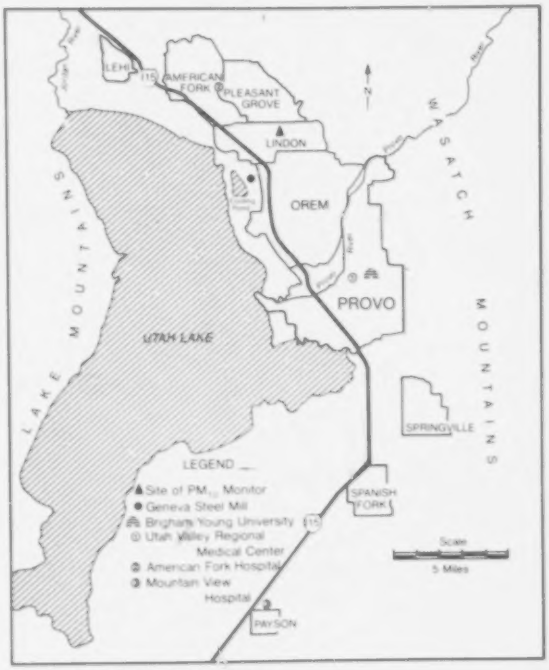

FIGURE 1—Study Area, Utah Valley

residents are members of the Church of Jesus Christ of Latter-Day Saints (Mormon)[25,26] which has strong church teachings against smoking.

Monitoring of particulate pollution began in 1964 and for carbon monoxide in 1971. On March 3, 1978, the EPA designated the county as a non-attainment area in accordance with provisions of Section 107 of the Clean Air Act. EPA ambient air quality standards for TSP and carbon monoxide were often exceeded at monitoring sites at Provo, Lindon, and Pleasant Grove during winter months when temperature inversions trapped emissions in stagnant air near the valley floor.

Generally, the county experienced improvements with respect to carbon monoxide pollution in the 1980s. At one monitoring site, the number of exceedances of the maximum eight-hour primary health standard for carbon monoxide fell from a high of 52 exceedances in 1982 to 10 exceedances in 1985. In order to continue to reduce levels of carbon monoxide pollution in the county, an automobile inspection/maintenance and anti-tampering program was implemented in 1986.

Particulate pollution levels in the county remained about the same from 1979–85. The 24-hour TSP standard of 260 $\mu g/m^3$ was exceeded as many as 10–18 times per year. The average annual geometric mean from 1979–85 for TSP at the Lindon monitor equalled 65 $\mu g/m^3$. This mean level of TSP exceeded EPA's annual secondary standard of 60 $\mu g/m^3$ but not the annual primary health standard of 75 $\mu g/m^3$. Monitoring of sulfur dioxides (SO_2) was conducted in the county in the 1970s but was discontinued because SO_2 levels were substantially below the annual primary health standard of .03 ppm, the 24-hour primary health standard of .14 ppm and the secondary 3-hour standard of .5 ppm.

The primary industrial source of fine particulate pollution as measured by PM_{10} in Utah County is the Geneva steel mill, commonly referred to as Geneva, located near Orem (Figure 1). When in operation, the mill emits approximately 82 per cent of all industrial sources of PM_{10} including power generation.[27] When all sources are accounted for, Geneva's contribution to PM_{10} equals 47 to 80 per cent of total emissions.[27] Other sources of PM_{10} include wood burning (approximately 16 per cent), road dust (approximately 11 per cent), diesel fuel and oil combustion (approximately 7 per cent). Also, Geneva's contribution to the county's industrial emissions of sulfur oxides, nitrogen oxides, hydrocarbons, and carbon monoxides are approximately 95, 98, 86, and 82 per cent, respectively.[27]

Geneva was built for the US Government in the early 1940s as part of the World War II effort. It was sold to US Steel Corp in 1946. On August 1, 1986, the Geneva steel mill shut down as a result of a labor dispute with USX Corporation (previously US Steel Corp.) The plant reopened on September 1, 1987 under a new owner, Basic Manufacturing and Technologies of Utah, Inc. In April 1985, the Bureau of Air Quality began to daily monitor PM_{10} at a site in Lindon (Figure 1) using a Sierra Anderson high volume sampler. During the winter season of 1985/86, Geneva was still open and 24-hour PM_{10} levels exceeded 150 $\mu g/m^3$ on 13 occasions. The highest single day concentration was 365 $\mu g/m^3$. During the winter of 1986/87 while Geneva was shut down, 24-hour PM_{10} levels never exceeded 150 $\mu g/m^3$. During the winter of 1987/88, following the reopening of Geneva, 24-hour PM_{10} levels exceeded 150 $\mu g/m^3$ on 10 occasions with a single day high at 223 (Figure 2).

During the winter of 1985/86, a random sample of county residents indicated that most residents thought that air quality was a serious problem and 29 per cent indicated that they had one or more members of their family who had health problems that were aggravated by air pollution.[28] During the winter of

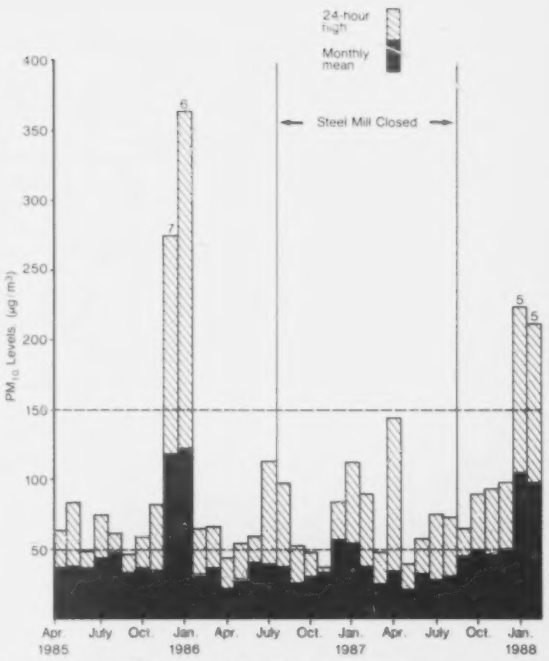

FIGURE 2—Monthly Mean and 24-Hour High PM_{10} (fine particulate pollution) Levels, Utah Valley, April 1985–January 1988

1987/88, following the closure and subsequent reopening of the steel mill, there was much local discussion about the contrast in air quality. The frequency and severity of respiratory illnesses were commonly perceived to have dropped when the mill was shut down, and then dramatically increased when it reopened. Newspaper articles, letters to the editor, and testimonials in public meetings often reflected this perception by many in the community.

Health Data

Hospital admissions data for respiratory-related illnesses were collected from April 1985 through February 1988. There were only four hospitals in the county. Data were collected from three of them that together had 579 beds. The other hospital in the County had only 20 beds, no pediatrics unit, no pulmonologist on its staff, and rarely provided inpatient care for respiratory illnesses.

A preliminary study of diagnosis-related groups (DRGs)[29] at Utah Valley Regional Medical Center indicated that DRGs 79, 80, and 81 (Respiratory Infections and Inflammations), DRGs 85 and 86 (Pleural Effusion), DRG 87 (Pulmonary Edema and Respiratory Failure), DRG 88 (Chronic Obstructive Pulmonary Disease), DRGs 92 and 93 (Interstitial Lung Disease), and DRGs 99 and 100 (Respiratory Signs and Symptoms) accounted for only 4.9, 0.6, 1.6, 3.3, 0.8, and 5.8 per cent of the collected cases, respectively. Initial comparative statistical analysis and regression analysis did not reveal any association between the closing and subsequent reopening of Geneva or PM_{10} levels and hospital admissions for any of these DRGs individually or collectively.

The bulk of the respiratory illness (83 per cent) were for the six DRGs that included 89, 90, and 91 (Simple Pneumonia and

Pleurisy) and 96, 97, and 98 (Bronchitis and Asthma) with 42 and 41 per cent of the cases, respectively. As a result, this analysis focuses on hospital admissions where the principal diagnosis was classified within one of these six DRGs.

Monthly admissions data for these six DRGs were compiled for each of the three hospitals. Records for outpatient and emergency admissions were not complete or consistent for the full time period. Therefore, only inpatient data were used in this analysis. Accurate records for Mountain View Hospital were available for the time periods April 1985 through September 1986 and January 1987 through February 1988. Accurate inpatient records for the other two hospitals were available from April 1985–February 1988.

Analysis Conducted

Utah Valley Regional Medical Center's admissions were sorted into in-county and out-of-county admissions. Both Utah Valley Regional Medical Center and American Fork Hospital are located within the central urban area of the county, near the major sources of pollution. The primary analysis used the combined Utah Valley Community Hospital in-county admissions and American Fork Hospital admissions as an indicator of the level of relatively severe respiratory illness in the urban area of the county.

Three other sets of hospital admissions data were used as control variables: "all-other" admissions from Utah Valley Regional Medical Center and American Fork Hospital, excluding in-county admissions for pneumonia, pleurisy, bronchitis and asthma; out-of-county admissions to Utah Valley Regional Medical Center for pneumonia, pleurisy, bronchitis, and asthma; and admissions to Mountain View Hospital in Payson for the same illnesses.

Hospital admission levels were compared across months with different levels of particulate pollution as measured by PM_{10}. Admission levels were also compared across comparable periods of time when the steel mill was open, closed, and then reopened. Finally, monthly hospital emissions were regressed on PM_{10} levels and weather variables obtained at Brigham Young University.[30]

Results

Comparative Analysis

As can be seen in Figure 2, there was considerable variability in PM_{10} levels in the county over the study period. During those months when exceedances of the 24-hour PM_{10} standard of 150 $\mu g/m^3$ occurred, the number of admissions for children, 0–17 years of age, was nearly triple the number of admissions for months with no exceedances (Table 1). In adults, admissions were approximately 44 per cent higher during the months when exceedances occurred.

During months when the arithmetic mean PM_{10} levels were equal to or greater than 50 $\mu g/m^3$, children admissions were nearly double than when the average PM_{10} levels were less than 50 $\mu g/m^3$. Adult admissions were increased by approximately 47 per cent.

The above comparisons were complicated by the fact that the months with especially high levels of particulate pollution were during the winter, and the reason for the high incidence of respiratory illness may be at least partly attributed to winter weather. The intermittent operation of the steel mill provided the opportunity to compare different winter seasons with marked differences in PM_{10} levels. Figure 2 demonstrates that when the steel mill was closed, PM_{10} levels were relatively low. One concern about making observations pertaining to these time periods is that the winter when the Geneva steel mill was closed may have had relatively good weather conditions and limited conditions of stagnant air. Weather data indicated that temperatures fell below zero on only two occasions throughout the study period, both in January of 1988 when they fell as low as $-1°$ and $-7°$. Snowfall during this time period for 1985/86, 1986/87, and 1987/88 totaled only 45.5, 33.5 and 27.5 inches, respectively.[30]

The National Weather Service computes an air stagnation or clearing index for valleys in Western Utah, including Utah Valley.[31] The index ranges from 0–1000 with lower values indicating more stagnant air. When the index is less than 200 pollution dispersal is "very poor" and weather conditions are such that air pollution potential is high. The month with the lowest average clearing index occurred during the winter the mill was closed. The average clearing index for the winter period of December–February for 1985/86, 1986/87, and 1987/88 was 388, 345, and 367, respectively, and the number of days when the index was below 200 for the same time periods equalled 47, 54, and 47, respectively.[32] Based on this index the air was relatively more stagnant and had higher air pollution potential during the winter when the mill was shut down than the previous or following winters.

Table 2 presents comparisons of hospital admissions be-

TABLE 1—Comparisons of Monthly Average Number of Hospital Inpatient Admissions for Utah Valley Regional Medical Center and American Fork Hospital across Months with Different Levels of PM_{10}[a]

Months Included	Number of Months Included	Mean PM_{10} Level for Months Included	Mean High PM_{10} Level for Months Included	Bronchitis and Asthma Ages 0–17	Bronchitis and Asthma Age 18+	Simple Pneumonia and Pleurisy Age 0–17	Simple Pneumonia and Pleurisy Age 18+	Subtotal Ages 0–17[b]	Subtotal Age 18+[b]	TOTAL[b]
All months	35	45.8	94.7	12.5	17.5	12.0	22.7	24.5	40.2	64.7
		(4.3)	(11.9)	(1.6)	(1.0)	(1.5)	(1.6)	(2.8)	(2.3)	(4.5)
Months when 24-hour $PM_{10} < 150$ $\mu g/m^3$	31	37.5	72.3	10.5	16.9	9.9	21.4	20.4	38.3	58.6
		(1.6)	(4.4)	(1.2)	(1.0)	(1.1)	(1.3)	(1.9)	(2.0)	(3.5)
Months when 24-hour $PM_{10} > 150$ $\mu g/m^3$	4	110.3	268.5	27.8	22.3	28.3	33.0	56.0	55.3	111.3
		(5.5)	(35.0)	(6.7)	(2.9)	(4.6)	(8.1)	(11.1)	(10.0)	(14.0)
Months when mean $PM_{10} < 50$ $\mu g/m^3$	27	35.1	68.7	10.1	16.5	10.2	19.8	20.3	36.3	56.7
		(1.3)	(4.6)	(1.3)	(1.1)	(1.2)	(1.2)	(2.1)	(2.0)	(3.9)
Months when mean $PM_{10} \geq 50$ $\mu g/m^3$	8	82.0	182.5	20.4	20.8	18.0	32.5	38.4	53.3	91.6
		(11.0)	(36.4)	(4.4)	(1.8)	(4.5)	(3.9)	(8.5)	(5.1)	(10.0)

[a]Standard errors at the means are presented in parentheses.
[b]Total may not sum up exactly due to rounding error.

tween fall and winter periods when the steel mill was open, closed, and reopened. During the winter months from December to February, hospital admissions for children were approximately three times as high when the steel mill was open than when it was closed. Even during the Fall months (September–November) when no exceedances for the 24-hour primary health standard occurred, children admissions for bronchitis and asthma were approximately twice as high when the steel mill was open. Adult hospital admissions were not as obviously associated with the reductions of PM_{10} that accompanied the closure of the steel mill. There was, however, a notable increase in adult admissions following the reopening of the mill.

Regression Analysis

The results of some of the regression models are presented in Table 3. Model 1 regresses total monthly hospital admissions for pneumonia, pleurisy, bronchitis, and asthma on current and lagged PM_{10} levels. All lagged variables simply refer to the previous month's value. The results demonstrate a strong correlation between admissions and PM_{10}. In fact, 59 per cent of the variance in monthly admissions for these respiratory illnesses is explained by current and lagged monthly mean PM_{10} levels alone.

In Model 2, current and lagged mean low temperature variables were also included. This relatively simple linear model with only PM_{10} and temperature variables explains 83 per cent of the variance in total hospital admissions for these respiratory illnesses. The correlation between mean PM_{10} levels, mean low temperatures and hospital admissions is particularly striking when actual admissions and estimated admissions based on Model 2 are plotted together over time (Figure 3). Models 3–14 repeat the analysis done in Models 1 and 2 for total adult admissions, total children admissions, adult admissions for pneumonia and pleurisy, children admissions for pneumonia and pleurisy, adult admissions for bronchitis and asthma, and children admissions for bronchitis and asthma.

Autocorrelated errors exist with some of the models, particularly those with only PM_{10} levels as independent variables. This autocorrelation, however, is largely eliminated when weather variables are included. For example, the Durbin-Watson D statistic is 1.0 for Model 1 and 1.6 for Model 2. It is 1.3 for Model 3 and 2.0 for Model 4. There is also collinearity between PM_{10} levels and temperature. The correlation coefficient between the mean low temperature and monthly mean PM_{10} levels equals −0.32. This collinearity complicates the analysis and makes specific best point estimators of the correlation coefficients difficult to estimate. However, Model 2 was reestimated using a nonlinear quasi-Newton iterative procedure which gave identical regression coefficients with somewhat smaller standard-errors.

Numerous other regression models were estimated that included snowfall, rainfall, evaporation, monthly mean temperatures, and mean high temperatures. The weather variable that was consistently most highly correlated with admissions was the mean low temperature. Regressions that used PM_{10} levels lagged for two months, and dummy variables that indicated the opening and closing of the steel mill and inversion seasons were also tried. Even with the inclusion of these other variables, strong, positive, correlations between hospital admissions and PM_{10} levels remained. Regression models were also estimated with monthly 24-hour high PM_{10} levels used as independent variables. The results were similar to those in Models 1–14 as presented in Table 3, but 24-hour high PM_{10} levels were generally not as strongly correlated with admissions as were the mean PM_{10} values.*

Analysis with Control Variables

Neither comparative analysis nor regression analysis revealed any associations between the control variables and PM_{10} levels or the closing and reopening of the steel mill. "All-other" admissions that excluded in-county admissions for pneumonia, pleurisy, bronchitis, and asthma averaged 1,562 per month and appeared to be declining slightly over the study period. No seasonal variability nor any association with PM_{10} levels or the closing and reopening of the mill was observed. Monthly "all-other" admissions regressed on PM_{10} levels and temperature variables (Models 15 and 16 in Table 3) showed no significant correlation with PM_{10} levels.

Out-of-county hospital admissions to Utah Valley Regional Medical Center and admissions to Mountain View Hospital in Payson were regressed on PM_{10} levels and temperature variables. Models 17 and 18 in Table 3 present the results of the regressions for total out-of-county admissions for pneumonia, pleurisy, bronchitis, and asthma. The same regressions were also run on out-of-county and Mountain View Hospital with admissions broken down by adults, children, and respiratory illnesses, as done in Models 1–14.

*Data available upon request to author.

TABLE 2—Comparisons of Hospital Inpatient Admissions for Utah Valley Regional Medical Center and American Fork Hospital across Time Periods with Geneva Steel Mill Open and Closed

Year	Steel Mill Open?	Mean PM_{10} Level for Months Included	Mean High PM_{10} Level for Months Included	Bronchitis and Asthma Ages 0–17	Bronchitis and Asthma Age 18+	Simple Pneumonia and Pleurisy Ages 0–17	Simple Pneumonia and Pleurisy Age 18+	Subtotal Ages 0–17	Subtotal Age 18+	TOTAL
				Winter Months (December–February)						
1985/86	yes	90	235	78	75	76	73	154	148	302
1986/87	no	51	96	23	67	32	83	55	150	205
1987/88	yes	84	177	78	65	71	126	149	191	340
				Fall Months (September–November)						
1985	yes	35	63	49	46	20	51	69	97	166
1986	no	31	47	23	48	25	60	48	108	156
1987	yes	47	83	55	46	24	66	79	112	191
				Fall and Winter (September–February)						
1985/86	yes	63	149	127	121	96	124	223	245	468
1986/87	no	41	71	46	115	57	143	103	258	361
1987/88	yes	66	130	133	111	95	192	228	303	531

TABLE 3—Sample Results of Multiple Regression Analysis

Model	Dependent Variable: Monthly Admissions	Regression Coefficients[a] Constant	PM_{10} Mean	Lagged PM_{10} Mean	Low Temperature	Lagged Low Temperature	R^2
1	Total	21.18	0.357	0.599	—	—	.59
		(7.1)	(.14)	(.15)			
2	Total	95.54	0.119	0.339	−0.351	−0.929	.83
		(12.8)	(.11)	(.11)	(.30)	(.31)	
3	Total Adult	25.31	0.150	0.175	—	—	.26
		(4.9)	(.09)	(.10)			
4	Total Adult	73.65	−0.016	0.017	−0.347	−0.486	.64
		(9.4)	(.08)	(.08)	(.22)	(.23)	
5	Total Child	−4.14	0.207	0.425	—	—	.67
		(4.0)	(.08)	(.08)			
6	Total Child	21.89	0.135	0.321	−0.004	−0.443	.75
		(9.7)	(.08)	(.08)	(.23)	(.24)	
7	Pn/Pl Adult	14.57	0.139	0.034	—	—	.19
		(3.5)	(.07)	(.07)			
8	Pn/Pl Adult	46.84	0.020	−0.063	−0.305	−0.252	.54
		(7.3)	(.06)	(.06)	(.17)	(.18)	
9	Pn/Pl Child	−1.50	0.112	0.183	—	—	.53
		(2.5)	(.05)	(.05)			
10	Pn/Pl Child	15.49	0.086	0.095	0.196	−0.487	.72
		(5.3)	(.04)	(.05)	(.13)	(.13)	
11	Br/As Adult	10.74	0.011	0.140	—	—	.36
		(2.0)	(.04)	(.04)			
12	Br/As Adult	26.81	−0.037	0.081	−0.042	−0.234	.59
		(4.3)	(.04)	(.04)	(.10)	(.11)	
13	Br/As Child	−2.63	0.094	−0.241	—	—	.60
		(2.5)	(.05)	(.05)			
14	Br/As Child	6.40	0.049	0.226	−0.201	0.044	.64
		(6.5)	(.05)	(.06)	(.15)	(.16)	
15	All-Other	1586	−0.050	−0.604	—	—	.02
		(46)	(.9)	(1.0)			
16	All-Other	1482	0.840	−0.798	5.904	−4.069	.15
		(120)	(1.0)	(1.0)	(2.8)	(2.9)	
17	Out-of-County Total	15.09	−0.047	−0.006	0.123	−0.264	.43
		(2.9)	(.02)	(.03)	(.07)	(.07)	
18	Mountain View Total	33.38	−0.013	−0.041	0.073	−0.474	.46
		(6.8)	(.05)	(.06)	(.16)	(.16)	

[a]The absolute value of the standard errors is provided in parentheses.
Pn/Pl = pneumonia/pleurisy
Br/As = bronchitis/asthma

Although Payson is located in the county and should be similarly influenced by contagious illness, it is over 32 kilometers from the major sources of pollution and should not be as influenced by monitored levels of PM_{10}.

The regressions using out-of-county and Mountain View Hospital admissions are limited by the fact that only about 15 per cent of Utah Valley Regional Medical Center admissions are out-of-county, and Mountain View Hospital's data are missing for the months of October, November, and December of 1986. The results indicated that there was significant correlation between the mean low temperature lagged variable similar to those in the earlier regressions. There was no positive correlation between out-of-county or Mountain View Hospital admissions and PM_{10} levels, however.

Discussion

The results indicated that hospital admissions for respiratory illnesses were strongly associated with PM_{10} levels. This association is much stronger for children than adults, and somewhat stronger for bronchitis and asthma than for pneumonia and pleurisy. These associations were particularly strong with monthly lagged variables suggesting that the health effects of particulate pollution are cumulative and that it takes time before they are manifested in inpatient hospital admissions data.

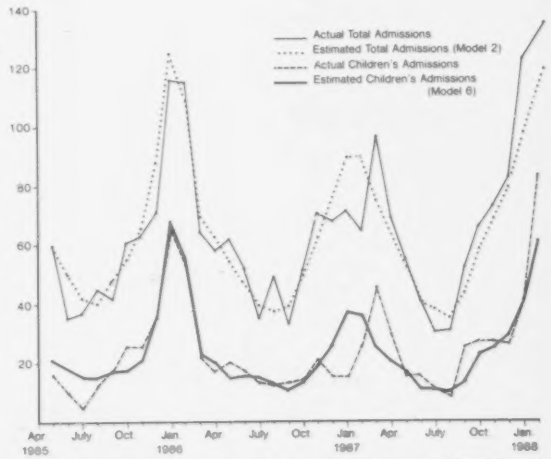

FIGURE 3—Actual and Estimated Hospital Admissions, April 1985 through January 1988, Utah Valley

Also, increased admissions for children are observed even for months when PM_{10} did not exceed 150 $\mu g/m^3$, suggesting that this standard may not be adequate protection for some children.

There are several concerns about these observations. One concern is that if increases in contagious illnesses such as influenza by chance coincided with periods of high PM_{10} levels, particularly during the winters when the steel mill was open, then the observed correlation between PM_{10} and admissions may be spurious. It would be expected, however, that if this were the case, the same correlation would be found in hospital admissions from neighboring counties or communities unaffected by Utah Valley's principal sources of pollution. No such correlation was found for out-of-county admissions to Utah Valley Regional Medical Center or to Mountain View Hospital in nearby Payson. Nor was such correlation found between PM_{10} levels and non-respiratory hospital admissions.

Another concern is that often levels of several air pollutants rise and fall in concert. PM_{10} may be a surrogate for other air pollutants with which it is temporarily associated. Two pollutants that may have had similar impacts on respiratory illnesses during the study period are total suspended particulates and ozone. Because PM_{10} measures only relatively small particles of particulate pollution, and because it is the smaller particles that are expected to pose the greatest health risks, it is considered the most appropriate measure of particulate pollution as it relates to respiratory health.[1] Regression models estimated with monthly mean total suspended particulate levels used as independent variables yielded results similar to Models 1–14 presented in Table 3 which used PM_{10}. The correlations between admissions and total suspended particulates were generally not as strong as those between admissions and PM_{10}.

There was no evidence that suggested that PM_{10} was serving as a surrogate for ozone pollution. The only times ozone pollution in Utah Valley rose to levels of any consequence was in the summer months during hot sunny days, whereas the periods of high levels of PM_{10} and hospital admissions for respiratory illness occurred mostly during the winter months when the steel mill was in operation. The results of this study suggest that the dominant pollution in terms of its impact on respiratory health in Utah Valley is particulate pollution and that PM_{10} is a better indicator than TSP.

Finally, the association between respiratory illness and particulate pollution found in this study is relatively large as compared with some previous studies.[20–22,33] This relatively strong association can be explained in part because PM_{10} is a better indicator of particulate pollution as it relates to respiratory health than previously used indicators.[1] Also, because Utah Valley experiences relatively high levels of particulate pollution, yet has an extremely low portion of its population that smoke, particulate pollution is likely a relatively large contributor to respiratory disease in the county.

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Public Health Briefs

Exposure to Occupational Hazards among Hispanics, Blacks, and Non-Hispanic Whites in California

James C. Robinson, PhD, MPH

Abstract: Relative risks of exposure to each of six types of occupational injuries and illnesses for Hispanic and Black workers compared to Whites who are not Hispanic were calculated using 1986 California data. Among males, Hispanics faced relative risks of exposure to all hazards adjusted for education and years of work experience of 1.33 (95% CI 1.22, 1.45), while Blacks faced relative risks of 1.17 (1.0, 1.37). Among females, adjusted relative risks were 1.19 (1.09, 1.29) for Hispanics and 1.31 (1.15, 1.50) for Blacks. (*Am J Public Health* 1989; 79:629–630.)

Introduction

Several studies have documented racial differences in exposure to occupational hazards, using national data on working conditions faced by Black and White workers.[1–4] These studies have been forced by the nature of the data they used to group Hispanics together with Whites who are not Hispanic. This approach not only conceals relative exposure differences between Hispanic and other Whites, it also may underestimate the difference in exposure levels between Blacks and Whites. The present study examined 1986 data on exposure to occupational injury and illness hazards for 5,737 workers in California, and calculated relative risks of exposure for Hispanics and Blacks, respectively, compared to Whites who are not Hispanic.

Methods

Information on work-related injuries and illnesses in California was obtained for 1986 from the mandatory reporting system of the state Workers' Compensation system. Under California law, employers must report to the state any work-related injury or illness that causes absence from work lasting one or more full days. These incidence reports are not directly influenced by whether or not the injuries and illnesses subsequently result in a compensation claim. The California Department of Industrial Relations tabulates these reports by occupation, industry, type of injury or illness, and a number of other criteria.[5]

Data were obtained for each of the five major types of injury and for acute illnesses as well as for total injuries and illnesses for each three digit Census occupation classification. Frequencies were converted to rates per 1,000 workers per year using occupation-specific California employment data from the 1980 Census of Population,[6] inflated by the 1980–86 rate of growth in total California employment.

The five major classifications of occupational injuries, which in 1986 accounted for 78.4 per cent of all reported injuries and illnesses, were: strains, sprains, dislocations, and hernias (43.3 per cent); cuts, lacerations, and punctures (13.4 per cent); fractures (10.0 per cent); crushing injuries, contusions, and bruises (8.0 per cent); and abrasions and scratches (3.7 per cent). Acute illnesses accounted for 7.5 per cent of all events.

Demographic data were obtained from the March 1986 Current Population Survey (CPS) of the US Department of Commerce's Bureau of the Census.[7] The March 1986 survey contained data on 6,175 employed Californians, 5,737 of whom were employed in occupations for which injury and illness data were available. Of 2,649 female workers, 676 were Hispanics, 1,798 other Whites and 175 Blacks. Of 3,088 male workers, 968 were Hispanic, 1,966 other Whites and 154 Black.

Educational attainment as recorded by CPS was categorized as: not graduated from high school, high school graduates, less than four years of college, and four or more years of college. A potential work experience measure was constructed by subtracting years of education plus five from years of age.[8] The square of this measure was also employed, in order to capture any nonlinear association between work experience and labor market rewards.

The occupational injury and illness rates were matched with the demographic data using three digit Census occupation codes. Each individual worker was thus ascribed the level of risk for each type of injury and for acute illness that was the average for his or her three digit occupation.

Injury rates were first transformed into logarithmic units and regressed on dichotomous variables indicating Hispanic and Black ethnicity, respectively, with other Whites who were not Hispanic composing the comparison category. Relative risks were calculated by taking the exponent of the coefficients on the ethnicity variables. Relative risks were adjusted by taking the exponent of the coefficients on the ethnicity variables in multivariate regressions that controlled for the four education and two work experience variables. The injury and illness rates were approximately lognormally distributed.

Results

Among men, Hispanics faced unadjusted relative risks for all injuries and illnesses combined of 2.21 (95% confidence interval = 2.04, 2.40), compared to other Whites; Blacks

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TABLE 1—Relative Risks of Exposure for Male Workers, Adjusted for Levels of Education and Potential Work Experience

Injuries/Illness	Hispanics (95% CI)	Blacks (95% CI)
All Injuries and Illnesses	1.33 (1.22, 1.45)	1.17 (1.00, 1.37)
Sprains and Strains	1.33 (1.22, 1.45)	1.22 (1.04, 1.44)
Cuts and Punctures	1.42 (1.27, 1.57)	1.03 (0.85, 1.26)
Fractures	1.27 (1.16, 1.38)	1.05 (0.89, 1.23)
Crushing Injuries and Contusions	1.42 (1.29, 1.56)	1.24 (1.04, 1.48)
Abrasions	1.48 (1.32, 1.66)	1.03 (0.83, 1.28)
Acute Illness	1.25 (1.16, 1.34)	1.19 (1.04, 1.36)

TABLE 2—Relative Risks of Exposure for Female Workers, Adjusted for Levels of Education and Potential Work Experience

Injuries/Illness	Hispanics (95% CI)	Blacks (95% CI)
All Injuries and Illnesses	1.19 (1.09, 1.29)	1.31 (1.15, 1.50)
Sprains and Strains	1.18 (1.08, 1.30)	1.40 (1.22, 1.62)
Cuts and Punctures	1.26 (1.13, 1.40)	1.12 (0.95, 1.33)
Fractures	1.11 (1.03, 1.19)	1.22 (1.08, 1.37)
Crushing Injuries and Contusions	1.24 (1.13, 1.36)	1.32 (1.14, 1.53)
Abrasions	1.30 (1.17, 1.44)	1.30 (1.11, 1.53)
Acute Illness	1.15 (1.07, 1.23)	1.30 (1.17, 1.45)

faced relative risks of 1.41 (95% CI = 1.18, 1.69). Hispanic men had unadjusted relative risks greater than those of Blacks in each individual category of injury and illness. Hispanic women faced relative risks of exposure to all occupational hazards of 1.49 (95% CI = 1.38, 1.62), while Black women faced relative risks of 1.31 (95% CI = 1.14, 1.50).

Tables 1 and 2 present relative exposure risks for Hispanics and Blacks compared to other Whites who were not Hispanic after controlling for educational level and potential work experience. The size of the relative exposure risks are reduced, but substantial ethnic differences remain. Among Hispanic workers, the relative risk for all exposures declined from 2.21 to 1.33 for males and from 1.49 to 1.19 for females. The relative risk for all exposures declined from 1.41 to 1.17 for male Black workers, but for female Blacks were almost identical to their unadjusted relative risks.

Discussion

Hispanic and Black workers in California are exposed to higher risks of occupational injury and acute illness than are other White workers who are not Hispanic, even after controlling for years of education and potential work experience. These ethnic differences may be associated with differing job characteristics. Hazardous occupations have been found to offer fewer rewards and amenities in most dimensions than non-hazardous occupations, including fewer opportunities for worker control over the pace and content of the work process; fewer possibilities for on-the-job training and promotion[9]; higher rates of temporary and permanent layoffs[10]; and lower wages.[9,11]

The limitations of these findings should be stressed. No data sources currently exist that include information on health and safety hazards in particular jobs and on the ethnic and socioeconomic characteristics of the workers employed in them. This study combined three different data sets: injuries and illnesses from state Workers' Compensation records, employment data from the Census of Population, and demographic and socioeconomic data from a large survey of individuals. The limitations are especially serious with respect to occupational illnesses as distinct from injuries, since Workers' Compensation programs are mainly focused on acute illnesses rather than diseases with long latency periods. Efforts to improve the quality of data on occupational illness and injury for surveillance purposes should include ethnic and socioeconomic information wherever possible.

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Lead Toxicity and Iron Deficiency in Utah Migrant Children

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Abstract: We determined the frequency of presumptive iron deficiency and lead toxicity in 198 Utah migrant children, ages 9–72 months, during the summer of 1985. There were no confirmed cases of lead toxicity. Thirteen per cent of those tested and 30 per cent of the children ages 9–23 months were iron deficient. Hematocrit determinations accurately predicted iron deficiency in only 35 per cent of the children confirmed to have this disorder via erythrocyte protoporphyrin screening. (*Am J Public Health* 1989; 79:631–633.)

Introduction

Lead toxicity and iron deficiency have their highest prevalence among children of low socioeconomic status.[1] Their hematologic and nonhematologic manifestations reflect impaired enzymatic and metabolic processes. Both conditions impair normal growth and development in children.[2–14]

The Centers for Disease Control (CDC) recently lowered the toxicity standards in children from a blood lead level of 1.45 μmol/L to 1.21 μmol/L.[15]

In some ethnic groups,the use of folk remedies has led to unintentional lead poisoning of children.[16] Sixteen cases of lead poisoning in Mexican American children due to the use of greta or azarcon were reported between 1981 and 1985.[17,18] This folk remedy is a fine powder whose content is 70 to 90 per cent inorganic lead and whose estimated lifetime usage among Mexican Americans in the American southwest is 10 per cent.[19]

Of the migrant children in Utah, approximately 80 per cent are Mexican American whose home base is either Mexico or the lower Rio Grande Valley in Texas.* The migrant children in Utah are representative of the population thought to be at risk for lead toxicity related to folk remedy ingestion.

The screening test used in our study, erythrocyte protoporphyrin (EP), enabled us to determine the frequency of both lead toxicity and iron deficiency in the children evaluated.[20–23]

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Methods

During the summer of 1985, 334 migrant children were enrolled in seven Head Start centers in Utah. Of these children, 269 were in the targeted age range of 9–72 months.[24] We tested 198 children (73.8 per cent of those eligible) for lead toxicity and iron deficiency after obtaining informed consent from their parents.

Three medical students performed the screening tests for lead toxicity and iron deficiency. They adhered to the protocol for micro blood sampling that was published by the CDC's Blood Lead Proficiency Testing Program.[25]

The EP and hematocrit samples were collected in 100 μl heparinized capillary tubes. Within two to six hours of the sampling, the hematocrit determinations were made using a Clay Adam's Triac Centrifuge (Model #0200). The EP tests were made with a front-faced hematofluorometer (ESA Zinc Protoporphyrin), which underwent factory standardization at the start of the study and was calibrated daily with control samples.

At the time of the capillary sampling, an additional two 100 μl lead-free (ESA) capillary tubes were collected from each subject for blood lead determinations. Blood lead analyses were performed at a CDC accredited laboratory using anodic stripping voltammetry (ESA Model 3010A). The precision of the testing method was confirmed by running dual capillary samples for blood lead determinations for the first 90 migrant children screened. The Pearson correlation (r) was 0.94 for these paired samples.

Children having EP levels less than 0.62 μmol/L were not tested further. If the EP were greater than .62 μmol/L and the lead level were less than 1.21 μmol/L, iron deficiency would be diagnosed according to the guidelines established by Yip in 1983.[26] These guidelines have a sensitivity of 88 per cent, a specificity of 90 per cent, and a positive predictive value of 38 per cent based on a population prevalence of 6.6 per cent.

Children with both elevated EP (≥.62 μmol/L) and lead (≥1.21 μmol/L) levels underwent venipuncture to confirm the elevated lead level. If confirmed, a diagnosis of lead toxicity would be made according to CDC standards.[27] Iron and total iron binding concentration levels (TIBC) were then run. An iron-to-TIBC ratio of 0.12 or less would confirm the concurrent diagnosis of iron deficiency.[28]

This study protocol was approved by the Institutional Review Board at the University of Utah.

Results

Table 1 summarizes the age, racial, and gender distribution of the migrant children screened. There were no confirmed cases of lead toxicity, and no cases of lead contamination in the capillary tube samples sent for analyses. Table 2 summarizes the frequency of iron deficiency according to age and racial categories. Of the 198 migrant children tested, 26 (13.1 per cent) were iron deficient according to our protocol; 29.7 per cent of the children ages 9–23 months were iron deficient compared to 13.5 per cent and 3.1 per cent of the children ages 24–47 months and 48–71 months.

TABLE 1—Age, Sex, and Racial Distribution of Children Screened

Characteristics	% Migrant (N:198)
Age	
9–23 months	18.7
24–47 months	48.5
48–71 months	32.8
Sex	
Male	44.9
Female	55.1
Race	
Caucasian	2.5
Mexican American	75.3
Native American Kickapoo	12.6
Native American Navajo	9.6

TABLE 2—Percentage of Migrant Children Iron Deficient According to Race and Age

Race	% 9–23 Months	% 24–47 Months	% 48–71 Months	Total
Hispanic	22.7% (22)	16.4 (73)	1.9 (54)	12.1 (149)
Native American Kickapoo	55.6 (9)	7.7 (13)	0.0 (3)	24.0 (25)
Native American Navajo	20.0 (5)	0.0 (8)	16.6 (6)	10.5 (19)
Caucasian	0.0 (1)	0.0 (2)	0.0 (2)	0.0 (5)
Total	29.7 (37)	13.5 (96)	3.1 (65)	

NOTE: Number screened is shown in parenthesis.

Hematocrit determinations were not strongly predictive of iron deficiency. Of 37 children age nine months to two years, only four had hematocrits less than 35 per cent, while 11 of the 37 were iron deficient based upon EP criteria.

Discussion

No cases of lead toxicity were confirmed in the population we studied. Although the study adhered to lead toxicity standards established by the CDC, this screening method has a sensitivity of only 50 per cent in detecting blood lead levels greater than 25 μg/dl.[30] Thus, it is possible that a few elevated blood lead levels may have been missed. Moreover, a recent statement on childhood lead poisoning by the American Academy of Pediatrics suggests that evidence is mounting to lower the level of blood lead which would be considered elevated.[31,32]

We have demonstrated that capillary sampling for lead determinations can be carried out for large numbers of children in remote settings with no breakage or contamination of specimens. This sampling technique was well tolerated by the children.

Iron deficiency was common among these migrant children, especially in the nine months to two year age group. This is the time period when the growth and development of children are most sensitive to the adverse effects of iron deficiency.[11–19] Migrant children should be included with other groups that have high prevalence of iron deficiency (urban Blacks, Indochinese refugees, and other low socioeconomic groups).[33,34]

We have shown, as have others, that the hematocrit determination is an insensitive screen for iron deficiency.**,[34] The office-based clinician generally has only this test available; thus, the problem of undetected iron deficiency remains a real one. The use of the EP as a screening test for iron deficiency seems warranted, particularly for those practitioners serving high-risk populations.

**Gloyd S, Meserole L, *et al*: unpublished data on screening for lead exposure and iron deficiency in southwest King County (Seattle) children, 1985.

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The Relationship of Park Maintenance and Accessibility to Dogs to the Presence of *Toxocara* spp. Ova in the Soil

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Abstract: Three parks in St. Joseph/Benton Harbor, Michigan were examined for the presence of *Toxocara* spp. ova in the soil. Nineteen per cent (22/114) of the samples contained ova, and mean density (eggs/5 g soil) was 0.73. The density of two samples, collected at the base of playground equipment in one park, was >2.1, which is considered sufficient for human infection. Subjective assessments of park maintenance and accessibility to dogs were not good indicators of the presence of *Toxocara* spp. ova. (*Am J Public Health* 1989; 79:633–634.)

Introduction

Visceral larva migrans (VLM) is a disease caused by the migration of the larval nematodes, normally parasitic in animals other than humans, through the extraintestinal tissues of humans.[1] *Toxocara canis*, the common roundworm of dogs, is regarded as the primary agent of this disease. Risk of human infection is greatest in young children, and most likely to occur via ingestion of embryonated eggs in the soil rather than through direct contact with an infected animal.[1]

The purpose of this study was to survey three parks in the St. Joseph/Benton Harbor area of southwestern Michigan for the presence of *Toxocara* spp. ova in the soil. Selection was based on subjectively determined differences in accessibility and levels of maintenance.

Methods

Study Sites

Milton Park, a well-kept neighborhood park used by children and adults, is located in an older neighborhood with some commercial development. It is approximately 70 m × 75 m, and is regularly maintained. There are conspicuous "No Dogs Allowed" signs posted around the park border.

Napier Park, a small triangular play area, approximately 45 m on a side, is located in a middle-class suburban neighborhood, frequently used by neighborhood children. Regular maintenance appears minimal, and there are no signs restricting access by dogs.

Corronde Park, the largest of the three, is bordered on one side by the St. Joseph River and regularly maintained. Facilities include: tennis courts, a baseball field, and picnic pavilions. It is isolated from surrounding housing and is typically used for picnics and family gatherings. Dogs are permitted but must be on a leash or chained. Soil samples were collected in a large play area (100 m × 55 m) adjacent to the river.

Soil Collection

Samples were collected at 9 m intervals, along transect lines approximately 9 m apart. Composite soil samples, totaling 40 g, were taken to a depth of 1–2 cm. Five to six plugs of soil were taken with a soil corer within a one square foot area at each grid point. Additional samples were collected at the bottoms of slides, bases of teeter-totters, and underneath swings. Samples were not collected when grid points fell on sidewalks or other obstructions.

Laboratory Procedures

Collection of eggs followed Method 6 of Kazacos[2] with minor variations, using 20 gram subsamples from each site. The specific gravity of the $NaNO_3$ used in the final flotation was 1.30 (Kazacos, personal communication). No attempt was made to differentiate between eggs of *Toxocara canis* and *T. cati*.

Results

Nineteen per cent (22/114) of the samples examined were positive for *Toxocara* spp. ova (Table 1). No ova were found in Milton Park, while Napier and Corronde Parks had similar prevalences of approximately 30 per cent. Mean density (ova/5g of soil) is higher in Corronde, but mean density for Napier falls within the 95% confidence limits of the differences between the two parks.

The five transects at Corronde Park fell into two distinct groups (Table 2). Two transects nearest the river (A & B) had only one positive sample between them. The remaining transects (C, D, and E) passed through the general play area, with prevalences between 50 and 60 per cent.

TABLE 1—Results of the Examination of Soil Samples for *Toxocara* spp. Ova from Parks in St. Joseph/Benton Harbor, Michigan

Parks	No. Samples	No. Positive	% Prevalence	Mean Density (range)
Milton	48	0	0	0
Napier	13	4	31	0.25
Corronde	53	18	34	0.83* (0.25,5.5)
Total	114	22	19	0.73 (0.25,5.5)

*Approximate 95% confidence interval for Corronde-Napier Parks is 0.08, 1.08.

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TABLE 2—Presence of *Toxocara* spp. Ova from five Transects* in Corronde Park

Transect	A	B	C	D	E
Ova present	0	1	6	5	5**
Ova absent	10	9	4	5	5

*Ten samples collected from each transect.
**Approximate 95% confidence interval for $Transect_{C,D,E}$–$Transect_{A,B}$ (48 per cent) is 26–68 per cent.

Discussion

Overall prevalence of *Toxocara* spp. ova from soil samples in this study (19 per cent) falls in the mid-range of other studies of public parks and playgrounds in North America.[3–6] Reasons for variation include, but are not restricted to, environmental conditions, choice of sampling sites, the utilization of the area by dogs and cats, and the prevalence of infection of *Toxocara* spp. in animals in the area.[7]

Subjective assessment of maintenance and access was of little value as a predictor of contamination. No ova were found in Milton Park, which is well maintained and has conspicuous signs prohibiting dogs. Napier and Corronde Parks differ markedly in appearance, maintenance, and accessibility, therefore, the similarity in prevalence (Table 1) must be based on the type of utilization. The majority of dogs in Corronde Park are brought by their owners, but must be leashed or chained. Visits are likely to be prolonged, resulting in a high probability that the animal will defecate on the premises. Napier Park, on the other hand, is more likely to be visited for shorter durations, e.g., by strays or pet owners taking animals there specifically for the purpose of defecation, rather than allowing the animal to defecate on private property and risk confrontation with the owner or resident.

Prevalence of *Toxocara* spp. ova is unquestionably a factor of how a park is actually used by pet owners rather than simply the accessibility of the area to animals. The primary factor is human behavior and not the behavior of the dog: strays would hardly be deterred by the presence of signs. Despite the fact that Napier Park appeared to have a lower level of maintenance and is more accessible to strays than Corronde Park, the mean density of ova was higher in Corronde.

The prevalence of eggs was higher in the three transects which pass through the general play area in Corronde Park than in the two transects immediately adjacent to the river. Two positive samples from Transect C (at the base of a large, jungle gym complex) had densities of 3.4 and 5.5/5 g of soil. These densities are above the level (2.1 ova/5g) considered sufficient to result in human infection,[8] although justification for this threshold value is obscure.

The numbers of eggs found in the present study were generally low, but this does not mean there is no cause for concern. If eggs are concentrated in areas where children play, the probability of infection is real even if the average for the park falls below the values considered necessary for human infection.

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American Indian Youth and Drugs, 1976–87: A Continuing Problem

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Abstract: Continuing surveillance of drug use among American Indian adolescents living on reservations shows them to have rates of use higher than those of their non-Indian counterparts. Marijuana use is particularly high among Indian students. By the 7th grade a significant number of Indian youth have tried drugs, particularly marijuana and alcohol, and there are few significant differences by gender. Based on observed patterns of use, intervention strategies need to begin in the elementary school years and target both males and females equally. (*Am J Public Health* 1989;79:634–636.)

Introduction

We have been collecting data on drug use rates among American Indian youth since 1975.[1–3] In addition to monitoring levels of use, we have conducted a series of studies examining the etiology of drug and alcohol use in this population.[4,5] In general, findings to date have shown that Indian youth have higher rates of drug use than non-Indian youth for nearly all drugs, and the interactions within small groups of peers (what we refer to as "peer clusters") have the strongest influence on a youth's decision to use or not use drugs. The latter finding is consistent with what we have

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TABLE 1—Per Cent of Indian 7th–12th Graders Reporting Lifetime Use of Drugs: 1975–87

Drugs	1975 N=1235	1977–78 N=3105	1980–81 N=2159	1982–83 N=1411	1984–85 N=1510	1986–87 N=2683
Alcohol	76	79	85	81	79	81
Get Drunk[a]	—	—	—	—	46	49
Marijuana	41	53	74	70	57	61
Inhalants	16	26	30	31	21	24
Cocaine	6	7	11	6	7	8
Stimulants[b]	10	15	24	22	21	25
Legal Stimulants[a]	—	—	—	—	14	15
Sedatives[b]	6	10	9	7	10	11
Heroin	3	4	5	2	5	5
Psychedelics	7	9	9	6	9	10
Tranquilizers[a,b]	—	9	6	3	7	7
PCP[a]	—	—	—	—	10	10
Cigarettes[a]	—	—	—	—	79	78
Smokeless Tobacco[a]	—	—	—	—	—	58

[a]Data not available for earlier years.
[b]Only illicit, or non-prescribed, use is included.

found for non-Indian youth.[6] There is no clear explanation for the higher rates of use for Indian youth although one can speculate that conditions of poverty, prejudice, and lack of opportunity found in many Indian communities create social conditions that encourage substance use. Liban and Smart, for example, found that Indian and non-Indian youth had similar rates of drug use when controlling for socioeconomic status.[7]

This article will update the previous data on trends in drug use of Indian youth who live on reservations.

Method

Instrument

The data are collected through self-administered, anonymous questionnaires given to students during a regular school period by the classroom teacher. The instrument has been developed specifically for use with Indian youth, and numerous revisions based on administration to over 30,000 Indian adolescents have led to a highly reliable and valid survey.*

Sample

A probability sample among Indian youth is extremely difficult to establish. Populations are remote, scattered, highly mobile, and often ill defined. As an alternative, reservations are classified on several demographic dimensions and an annual sample is constructed to represent the range on those dimensions. The parameters considered in selection include reservation size, geographic location, degree of isolation, and cultural diversity. The latter dimension is difficult to define since there are over 400 distinct tribes in the United States. There are cultural clusters, however, that may include many tribes with similar traditions, language, and history. The sampling plan draws from different clusters in any given year. Five to six tribes are selected each year yielding a sample of 1,000 to 2,000 youth.

The results are aggregated into biennial reporting periods which reduces any sampling bias that may occur in the annual selection procedure. Once a tribe is selected the survey is administered to every youth in the tribe who is attending school. The exception is in the large tribes where a sample of schools is used. Absentees and dropouts are not included.

The data collection methodology has remained consistent since 1975 assuring comparability across years.

Results and Discussion

Lifetime Prevalence: 7th–12th Grades

The most general measure of drug use is "lifetime prevalence," the question "Have you *ever* used?" (Table 1). While this does not provide information on amount or intensity of current use, it does yield an indication of the level of exposure to a drug in a population. If lifetime prevalence is high, the drug is clearly available and accessible.

The more commonly used drugs (alcohol, marijuana, and inhalants) increased in use quite sharply between 1975 and 1981. A similar pattern is seen for many of the other drugs. The interval from 1981 and 1985 was one of a general decline in lifetime prevalence for the most commonly used drugs. Marijuana use is illustrative, with a decrease from 74 per cent to 57 per cent. Inspection of the last two reporting periods suggests no further decline. Despite a slight rise for some drugs in 1988, it is premature to interpret this as the beginning of another period of increase.

The pattern showing a rapid increase in use of the most commonly used drugs through about 1981 and a modest decline after that is shared by both Indian and non-Indian youth.[8] This would suggest that young people of both races are responding similarly to the broad scale societal forces that influence drug use.

Gender and Age

The best available data for non-Indian youth show higher prevalence rates of both lifetime and last 30 days drug use for males for nearly all drugs.[9] When the Indian data are examined, however, females use drugs at the same rate as males. The exceptions are lower lifetime prevalence rates for females for cocaine (6.1 vs per 9.8 per cent) and smokeless tobacco (45.8 vs 71.4 per cent) and higher rates for cigarettes (82.0 vs 74.6 per cent). It is evident that interventions must target both sexes equally.

A significant amount of drug use occurs among Indian 7th graders, indicating that prevention must begin in the

*Oetting ER, Beauvais F, Edwards R, Waters M: The clinical drug abuse scale: Book II, instrument development, reliability and validity. Fort Collins, CO, Rocky Mountain Behavioral Science Institute, 1984.

TABLE 2—Per Cent of Indian and Non-Indian High School Seniors Who Have Ever Tried Drugs and Who Have Used Them in the Past Month: 1986–87

	Ever Tried		Used in Last Month	
	Indian	Non-Indian[a]	Indian	Non-Indian[a]
Drugs	N=325	N=15,200	N=325	N=15,200
Alcohol	90.8	91.3	58.5	65.3
Get Drunk	69.8	—	38.7	—
Marijuana	77.5	50.9	36.5	23.4
Inhalants	16.6	15.9	1.8	2.5
Cocaine	12.3	16.9	3.7	6.2
Stimulants	37.2	23.4	9.1	5.5
Legal Stimulants	17.5	—	2.1	—
Sedatives	11.5	8.4	3.4	1.8
Heroin	5.5	1.1	0.3	0.2
Psychedelics	11.7	9.7	0.6	2.5
Tranquilizers	8.1	10.9	1.2	2.1
PCP	8.6	4.8	0.6	1.3
Cigarettes	80.3	67.6	38.3	29.6
Smokeless Tobacco	56.8	—	31.4	—

[a]National High School Senior Survey[9]

elementary school years. For example, by the time they are in the 7th grade, 28 per cent of Indian youth report at least one episode of getting drunk, 44 per cent have tried marijuana, 22 per cent inhalants, 12 per cent stimulants, and 72 per cent cigarettes.

Comparison with Non-Indian Youth

Table 2 shows the 1986–87 comparison between Indian and non-Indian high school seniors[9] for lifetime and last 30 days prevalence. The Indian seniors have higher rates for six drugs but lower rates for only one drug—cocaine. Note in particular that on both prevalence measures, Indian students are showing considerably higher rates of marijuana use.

Since very high dropout rates have been reported for Indian students, it is likely that the rates reported for them in Table 2 would be even more elevated if all potential high school seniors were surveyed.[10]

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Family Physicians' Views of Chiropractors: Hostile or Hospitable?

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Abstract: Family physicians in the State of Washington were surveyed about their knowledge and views about chiropractors; 79 per cent (476) responded. Sixty-six per cent indicated discomfort with what they believed chiropractors do while acknowledging their effectiveness for some patients; 25 per cent viewed chiropractors as an excellent source of care for some musculoskeletal problems and only 3 per cent dismissed chiropractors as quacks that patients should avoid; 57 per cent admitted having encouraged patients to see a chiropractor. These views are less negative than those of organized medicine. (*Am J Public Health* 1989; 79:636–637.)

Introduction

Chiropractic began 100 years ago when its founder, a grocer and "magnetic healer," discovered what he believed to be the universal cause of all disease (vertebral subluxations) and its cure (spinal manipulation). After a shaky start, chiropractic is now recognized in all 50 states and is reimbursed under Medicare, Medicaid and by private insurers.[1]

The American Medical Association (AMA) has denounced chiropractic as "quackery and cultism" for decades, and in 1965 declared it unethical for physicians to have any professional association with chiropractors. Organized medicine's hostility toward chiropractic has left the public and many physicians[2] with the impression that physicians in general believe that chiropractors are quacks. This study describes practicing physicians' actual views of chiropractors.

Methods

All 181 family physicians employed by Group Health Cooperative of Puget Sound (GHC) and a 50 per cent random sample of the 847 non-GHC members of the Washington Academy of Family Physicians were surveyed about their management of low back pain. Approximately 75 per cent of family physicians in Washington are members of the

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Increasing Response Rates in Physicians' Mail Surveys: An Experimental Study

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Abstract: It is becoming increasingly difficult to obtain high response rates in physicians' mail surveys. In 1983–84, we tested the effectiveness of two techniques among 604 Quebec physicians who had not responded to an initial letter. A handwritten thank you note at the bottom of the letter accompanying the questionnaire and a more personalized mailout package increased response rates by 40.7 per cent and 53.1 per cent, respectively, compared to control groups. (*Am J Public Health* 1989; 79:638–639.)

Introduction

Epidemiologists, health administrators, and other public health professionals often rely on mailed questionnaires to obtain data from physicians. Mailed questionnaires cost less than interviews, are more convenient for the physicians, and allow less socially desirable responses.[1–4] The major disadvantage is the low response rate with the attendant problem of non-response bias.[5–7]

Many variables have been shown to influence physician response to mailed surveys. These variables include questionnaire length and format,[8–9] postage,[9] characteristics of investigator,[10] follow-up mailings,[11] and rewards.[12] In the study reported here, we experimentally tested the effectiveness of two relatively inexpensive procedures on physicians' response rates to a follow-up letter: 1) a handwritten postscriptum on the cover letter thanking the physician personally for participating in the study; and 2) a more personalized mailout package. Because of the small amount of time physicians have to answer the huge volume of mail they receive, we hypothesized that the two techniques would be more likely to catch physicians' interest and result in better response rates.

Methods

The experiment was conducted during a 1983–84 mail survey of a stratified sample of Quebec physicians (729 general practitioners and 379 specialists) aimed at determining the level of physicians' support for a number of patient care issues. The 70-item closed-ended questionnaire took about 20 minutes to complete. The data collection took place over six months and included four mailings, two of which served for the experimental study reported here.

At the first mailing, all physicians received a cover letter, the questionnaire, and a postage-permit return envelope sent first class in an outgoing metered envelope. We addressed the envelopes with a computerized self-adhesive address label with the last name first.

Nonrespondents to the first mailing were randomized into two groups. About half of the subjects received with the questionnaire a follow-up letter with a handwritten postscript which read as follows "Thank you, Doctor X, for your participation in the study." The other half received the same letter but no handwritten thank you note in the post-scriptum.

Nonrespondents to the second mailing were again randomized into two groups. One group was sent a mailout package which, instead of using the computerized address label with the last name first, had the physician's name and address individually typed onto the envelope with title preceding name (e.g., "Doctor John Smith" instead of "Smith, John"); 2) the outgoing envelope was identified to the university rather than to the survey title; 3) the outgoing envelope was handstamped rather than metered; and 4) the postage permit return envelope was replaced by a handstamped return envelope. The control group received the same mailout package as the one used earlier. As suggested by Dillman,[13] the data collection process was completed by a fourth mailing.

Results

Both techniques were effective in increasing physicians' response rates (Table 1). The handwritten post-script increased response rates by 40.7 per cent since 30.4 per cent of physicians in the handwritten post-scriptum treatment group responded compared to 21.6 per cent in the control group. The personalized mailout package proved even more effective: 39.5 per cent of physicians who were sent a personalized mailout package responded compared to 25.8 per cent for the control group. This represents an increase of 53.1 per cent.

Although specialists did not respond as well as general practitioners, they seemed somewhat more sensitive to the experimental variables (Table 1). In both experiments, the difference in response rates between the experimental and control groups was somewhat higher for specialists than for generalists.

Discussion

Achieving high response rates in physicians' mail surveys is often a challenge. Rates below 70 per cent are often reported in the literature.[14–17] In this study, we were able to improve physicians' response rates by using two relatively inexpensive techniques of personalizing the mailed questionnaire. The handwritten thank you note at the bottom of the letter accompanying the questionnaire increased physicians' returns by 40.7 per cent whereas the personalization of the mailout package proved even more effective, increasing returns by 53.1 per cent.

The techniques were more effective among physicians less interested in the survey topic (the specialists). Although one cannot exclude a regression to the mean effect, the results may suggest that specialists were particularly sensitive to efforts made to gain their collaboration.

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TABLE 1—Response Rates after Each Mailing: Physicians' Survey, Quebec, 1983–84

		Generalists				Specialists				Total			
Mailing	Experiment	n	% RR	% CRR	Difference (95% CI)	n	% RR	% CRR	Difference (95% CI)	n	% RR	% CRR	Difference (95% CI)
First	No	729	51.2	51.2		379	34.6	34.6		1108	45.5	45.5	
Second	Yes	356	28.9	65.3		248	21.4	48.6		604	25.8	59.6	
	PS⁺	170	32.9		7.6	119	26.9		10.6	289	30.4		8.8
	PS⁻	186	25.3		(−1.8, 17.0)	129	16.3		(0.4, 20.8)	315	21.6		(1.8, 15.8)
Third	Yes	253	39.5	79.0		195	23.6	60.7		448	32.6	72.7	
	PER⁺	126	46.0		12.9	97	30.6		14.6	223	39.5		13.7
	PER⁻	127	33.1		(1.0, 24.8)	98	16.3		(2.9, 26.3)	225	25.8		(5.1, 22.3)
Fourth	No	153	24.2	84.1		149	21.5	69.1		302	22.8	78.9	

RR: Response rate
CRR: Cumulative response rate
PS⁺: Group exposed to the handwritten post-scriptum
PS⁻: Control group
PER⁺: Group exposed to the personalized mailout package
PER⁻: Control group

Even though personalization was undoubtedly effective in increasing physicians' returns, it is unclear why it worked. It is likely that it plays at different levels by increasing the probabilities that physicians will open the mailout, read the cover letter, and accept to complete the questionnaire. As others have emphasized,[18,19] more research is needed to understand the principles underlying people's behavior in relation to mailed questionnaires and to uncover significant interaction among survey techniques, survey content, and population.

It has been suggested that personalization in mail surveys can be counterproductive under certain circumstances, especially when it raises doubts in subjects' minds about the anonymity of their responses.[20,21] This study shows that it can be very effective if used appropriately. In physicians' mail survey, personalization is a technique that should be considered, particularly in follow-up mailings.

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How Free Care Improved Vision in the Health Insurance Experiment

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Abstract: We studied reasons for the improvement in the functional vision of enrollees receiving free care in the Rand Health Insurance Experiment. Among low income enrollees, 78 per cent on the free plan and 59 per cent on the cost-sharing plans had an eye examination; the proportions of those obtaining lenses were 30 per cent and 20 per cent, respectively. Visual acuity outcomes of low income vs non-poor enrollees were more adversely affected by enrollment in cost-sharing plans. Free care resulted in improved vision by increasing the frequency of eye examinations and lens purchases. (*Am J Public Health* 1989; 79:640–642.)

Introduction

Over half of the US population wears corrective lenses.[1] In 1972, Americans spent $352.2 million on visits to ophthalmologists[2] and, in 1977, $1.6 billion on optical goods.[3] In 1975, 20 per cent of Americans had out-of-pocket expenditures for optical care.[2] Among automobile drivers, vision impairments are a major cause of accidents.[4]

The Rand Health Insurance Experiment (HIE) demonstrated that cost sharing reduced the use of health services,[5,6] but participants receiving free care had better visual acuity at the experiment's end than did those randomized to cost-sharing plans.[7]

In this paper, we explain why enrollees receiving free care had better corrected visual acuity and describe the amount of care for correcting visual acuity that was received by a non-aged general population.

Methods

Health Insurance Experiment (HIE)

The Rand HIE, a randomized trial of cost sharing, was conducted between November 1974 and January 1982 in six cities, among 3,958 enrolled persons ages 14–61 in 2,005 families; 70 per cent were enrolled for three years and 30 per cent participated for five years. Details about study design, sites, and participant exclusions have been previously reported.[8,9] The sample was a random sample of six communities. With the important exception of those over the age of 61 at enrollment, the sample is otherwise reasonably representative of the United States population with respect to age, sex, income, education, proportion Black, and marital status. The age distribution of participants in the free and cost- sharing groups were equivalent: 13 per cent aged 14–17, 17 per cent aged 18–24, 30 per cent aged 25–34, 19 per cent aged 35–44, 13 per cent aged 45–54, and 8 per cent aged 55–61. Families were assigned to one of 14 cost-sharing plans that varied in the degree of cost-sharing. The primary variation was in the coinsurance rate: 0, 25, 50, or 95 per cent, but out-of-pocket expenses were limited to $1,000 per family per year with a reduced amount for low-income families. All plans covered ambulatory and hospital care, preventive services, dental services, mental health care, and prescription drugs. Vision services were covered, subject to varying cost sharing, as follows:

- One eye examination for refractive purposes per year;
- One pair of corrective lenses per year. There was an additional out-of-pocket cost for contact lenses, and sunglasses were not covered unless medically necessary;
- One pair of frames every two years, with a maximum payment based on the normal price of standard frames in that area. The coinsurance rate for these services was based on the coinsurance rate for a given plan. For example, if the coinsurance rate was 25 per cent, then 75 per cent of the charge for refraction, glasses, and frames was reimbursed.

This analysis considers all enrollees age 14–61 at enrollment who entered and remained enrolled for the entire duration of the study. It excludes enrollees who died during the study and those enrolled in the HMO (health maintenance organization) portion of the experiment.

Assessment of Vision Status

Enrollees completed a medical history questionnaire both at entry to and exit from the experiment. This included questions about use of corrective lenses, presence of vision problems in daily life (e.g., Can you read the newspaper? Can you recognize a friend across the street?), and history of vision care.

In addition, the proportion of enrollees with corrective lenses on entry was equivalent in the free and cost-sharing groups. A random 60 per cent of enrollees at entry and all enrollees at exit received vision tests with and without lenses. Each eye was tested separately with the Snellen Eye Chart or the illiterate E chart and with the Rosenbaum near vision card. Visual acuity when tested without corrective lenses was termed "natural" acuity; when tested with lenses "functional" acuity. Examinees who did not wear glasses and had natural far vision worse than 20/20 or who wore glasses and had functional far vision worse than 20/20 were tested for pinhole acuity.

Explicit quality-of-care criteria specific to vision problems were developed by a panel of physicians and were reviewed by ophthalmologists. These focused on both processes and outcomes of care, including the presence of impaired functional vision at the end of the experiment, and are presented in full elsewhere.[10]

Analysis

We constructed a tree (Figure 1) that considered the points at which vision could have been affected by free care. Each branch represents a path to normal or impaired vision. We defined our population of interest as the 2,399 enrollees with natural vision impairment at entry or exit. (This number differs slightly from that reported in Brook, *et al*,[8] because we considered enrollees with near or far vision impairment at

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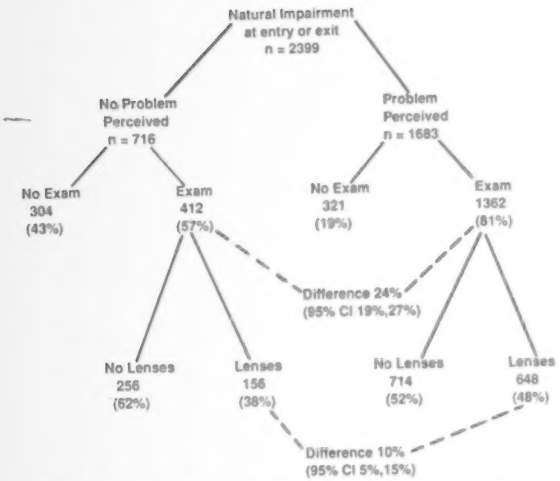

FIGURE 1—Relation of Perceived Vision Problem to Likelihood of Obtaining Lenses
All percentages refer to the n on the preceding branch.
Exam refers to eye examination during Health Insurance Experiment.

entry *or* exit). Because virtually no one in the experiment underwent a procedure to improve natural vision, such as cataract removal, we can assume that those with natural impairment at the beginning also had natural impairment at the end of the experiment and that the presence of such impairment was not affected by plan.

In this analysis, all cost sharing plans were grouped together into a single category, because the differences among cost-sharing plans with respect to vision outcomes were negligible.[9] Two-tailed T-tests were used to contrast variables on the free and cost-sharing plans.

Results

The prevalence of perceived vision problems at enrollment was 70 per cent on both free and cost-sharing plans. Those enrollees who perceived a vision problem were more likely to have an eye examination than those who did not, and those who both perceived problems and had their eyes examined were more likely to obtain lenses than those who did not perceive a problem but nonetheless received an examination (Figure 1).

TABLE 1—Effect of Cost-Sharing on Process of Care for Vision Problems

Measure	% on Free Plan (n)	% on Cost-sharing Plan (n)
Perceived vision problem	70 (579)	70 (1104)
Had eye examination if problem perceived	90 (519)	76* (843)
Purchased lenses if problem perceived and had eye examination	48 (250)	47 (398)
Purchased lenses if problem not perceived but had eye examination	41 (67)	36** (89)

*Difference 14% (95% CI 6, 22)
**Difference 5% (95% CI .1, 9.9)

Effect of Cost-Sharing on Care

Of those who perceived problems, 90 per cent on the free plan and 76 per cent on cost-sharing plans received an eye examination (Table 1). However, once eye examinations were performed on those who perceived vision problems, enrollees on free and cost-sharing plans obtained lenses with similar frequency. In contrast, enrollees who did not perceive a problem with their vision but who nonetheless had an eye examination were somewhat more likely to obtain lenses if they were on the free plan.

Poor enrollees, defined as those in the lower one-third of the income distribution, were as likely as non-poor enrollees to perceive problems with their vision (70 per cent and 71 per cent, respectively). Poor enrollees with natural impairment who were on cost-sharing plans were less likely to have an eye examination than those on the free plan. Moreover, conditional on having an examination, they purchased fewer lenses (Table 2). While these differences were also present for the non-poor, they were less pronounced.

We also studied those enrollees with functional near visual acuity worse than 20/40 at enrollment. Of this group, 8 per cent of the poor and 25 per cent of the non-poor on cost-sharing plans purchased lenses necessary to correct vision to 20/20, while there were no statistically significant

TABLE 2—Effect of Cost-Sharing on Vision Care for Poor and Non-poor Enrollees

	Free			Cost-sharing			Free Minus Cost-sharing	
	Poor (n)	Non-Poor (n)	Diff (95% CI)	Poor (n)	Non-Poor (n)	Diff (95% CI)	Poor (95% CI)	Non-Poor (95% CI)
Process								
% with natural impairment having an eye examination	78 (255)	86 (466)	8 (3,3)	59 (285)	74 (808)	15 (9,21)	19 (12,26)	12 (10,14)
% with exam and with natural impairment who purchased lenses	30 (82)	43 (235)	13 (6,20)	20 (98)	35 (389)	15 (11,19)	10 (4,16)	8 (3,13)
Outcome								
% who obtained lenses during the HIE if functional near vision at enrollment was 20/40 or worse in one or both eyes	18 (49)	22 (76)	4 (−3,9)	8 (74)	25 (147)	17 (13,21)	10 (5,15)	−3 (−9,3)
% with exit functional far visual acuity equal to or better than pinhole acuity	80 (340)	83 (622)	3 (−2,8)	76 (594)	81 (1275)	5 (2,8)	4 (0,−8)	2 (−1,4)

differences in the purchase of lenses between poor and non-poor on the free plan (Table 2).

At exit, somewhat fewer poor than non-poor enrollees on the cost-sharing plans had their vision maximally corrected, while the difference in visual acuity between the poor and non-poor groups on the free plan was a bit less (Table 2).

We compared results (i.e., whether vision was corrected) by plan type for enrollees who received a vision examination at entry and those who did not. Unlike the results for a similar analysis in hypertensives,[11] this demonstrated that differences in the screening entry examination did not improve the exit vision outcomes.

Table 3 shows the mean number of visits to an ophthalmologist or optometrist during the three years of the experiment. Both visits and lens purchases were more frequent for enrollees on the free care plan. In addition, high income enrollees received more vision care than did low income enrollees.

Discussion

It was previously reported that free care resulted in improved vision. One possibility was that participants on the free plan obtained corrective lenses more often than those on the cost-sharing plans. This was true for low-income enrollees. Our analysis suggests that a different mechanism accounted for improved vision in the general population—more eye examinations. It is also noteworthy that visual acuity outcomes for low-income enrollees were adversely and differentially affected by cost-sharing.

In contrast to hypertension,[11] there was no favorable effect of the entry examination on vision outcome in the cost-sharing group. This may be due to the fact that entry examination results were reported only to the enrollees' primary physician, not to their optometrist or ophthalmologist. The enrollee's primary physician may have been less likely to act upon a report of vision results than upon a report of poor blood pressure control.

The explanation that more frequent visits led to improved outcome held for three conditions: vision impairment, hypertension, and dental problems. All three of these conditions are relatively common and inexpensive to diagnose and treat. It is probable that the increased visit rate on the free care plan resulted in increased detection of these diseases. This effect was most pronounced for low income enrollees.

TABLE 3—Use of Services for Vision in the HIE Sample Over a Three-Year Period

	Poor			Non-poor		
	Free (277)	Cost-sharing (479)	Diff.	Free (544)	Cost-sharing (1099)	Diff.
Mean # visits (including 0)	1.63	0.92	.71	1.80	1.37	.43
Mean # lenses (including 0)	0.46	0.29	.17	0.69	0.50	.19

(sample sizes appear in parentheses)

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Financing Disposal of Non-municipal Wastes: A National Seminar

The Energy Bureau has announced a one-day seminar entitled "Financing Disposal of Non-Municipal Wastes"—an intensive program focusing on such wastes as medical and hazardous wastes, as well as municipal and industrial sludge. To be held at the Radisson Park Terrace Hotel in Washington, DC, on May 18, the seminar leaders will discuss and analyze the changing regulatory environment applicable to these wastes; their liability framework; financial issues for lenders, for real property acquisition, and for transportation, storage and disposal; as well as financing models, such as public and not-for-profit issuers, project finances, pooled financings and public/private partnerships; key risk management issues, such as insurance, change of law, site conditions, and waste stream control; and technological responses, such as incineration, fluidized bed construction, bioremediation and recycling. For further information, contact: Robert W. Nash, president, The Energy Bureau, 1133 Broadway, New York, NY 10010. Tel: (212) 633-8011.

Reflections on Curative Health Care in Nicaragua

R. Giuseppi Slater, MD

Abstract: Improved health care in Nicaragua is a major priority of the Sandinista revolution; it has been pursued by major reforms of the national health care system, something few developing countries have attempted. In addition to its internationally recognized advances in public health, considerable progress has been made in health care delivery by expanding curative medical services through training more personnel and building more facilities to fulfill a commitment to free universal health coverage. The very uneven quality of medical care is the leading problem facing curative medicine now. Underlying factors include the difficulty of adequately training the greatly increased number of new physicians. Misdiagnosis and mismanagement continue to be major problems. The curative medical system is not well coordinated with the preventive sector. Recent innovations include initiation of a "medicina integral" residency, similar to family practice. Despite its inadequacies and the handicaps of war and poverty, the Nicaraguan curative medical system has made important progress. (*Am J Public Health* 1989; 79: 646–651.)

Introduction

Few developing countries have attempted fundamental revision of their health care delivery systems to enable them to truly serve the poor who form the vast majority of their population. Since its victory over the Somoza dictatorship in 1979, the Sandinista revolution in Nicaragua has placed great emphasis on improving health care and has received international acclaim for its achievements in public and preventive health.[1–3] After nine years, it is also important to see how well the present medical system meets the never-ending demand for curative medical care. A realistic examination of the Nicaraguan experience is important to better understand the difficulties and possibilities inherent in seeking such change.

This article is based on my personal experience working for a year in the regional hospital and in a neighborhood health center in the northern city of Estelí, as well as several months' work with the new Nicaraguan family practice residency program in Managua. It is also based on extensive visits to many other health facilities in Nicaragua and numerous conversations with health personnel, both Nicaraguan and foreign, at all levels of the health care system. To the maximum extent possible, I have attempted to supplement personal observations with reference to the limited data and published material available about Nicaraguan health care.

Three broad criteria can be used to assess health care delivery: access and affordability, efficacy in diagnosis and treatment, and effectiveness in coordinating with the preventive health sector. In simpler terms, how easy is it to get health care, how good is the health care, and how much does the health care seek to maintain health as well as restore it? Viewed from these perspectives, Nicaraguan health care is a mixture of successes and problems.

Access and Affordability

Before the revolution, only a limited number of Nicaraguans (primarily public employees and National Guard members and their dependents) had guaranteed access to government-supported health care.[3–6] For most of the population, obtaining health care meant cash-up-front. People in Estelí described to me how the hospital, in Somoza's time, had a three-tier admission system: those with a certain amount of money were admitted to large open wards and received limited therapy; those with more funds were placed in semi-private rooms; and those with ample resources had private rooms and a readily available physician. Hospital workers reported watching patients die at the emergency entrance because they lacked the admission fee.

Since the institution of the Unified National Health System (SNUS) three weeks after the victory of the Sandinista Revolution in 1979, health care in public facilities has been offered to all Nicaraguans at no cost.[2,3] To fulfill this commitment, the Nicaraguan Government first sought to train more health workers and build more facilities. In the first several years after the revolution, the medical school curriculum was revised to emphasize primary care issues.[7] A second medical school was opened, in Managua, and class size was expanded in the original medical school in León, resulting in a four-fold increase in medical school enrollment between 1979 and 1984.[3,7] By 1984, residencies had been initiated in 16 medical specialties and master's degrees were offered in Epidemiology, Public Health, and Hygiene; previously, no postgraduate medical training had been available in Nicaragua. Nursing schools were expanded and a new program was begun to train laboratory technicians.[6,8,9] The number of primary care health facilities more than tripled over one decade, increasing from 172 in 1977 to 472 in 1984 and to 588 in 1987.[6,10]

The impact of the contra war, including the resultant economic disruption, has slowed or reversed some earlier gains,[2,11] as has continued rapid population growth which currently averages 3.3% per cent per year despite significant emigration.[12] Figure 1 illustrates the initial marked increase in the quantity of health care providers per 100,000 population in Nicaragua after the revolution as well as the subsequent leveling off; Figures 2 and 3 show similar progress in increased utilization of health care services as well as the recent declines. Reflecting overall progress, including public health advances, infant mortality dropped from 121/1000 to 76/1000 live births between 1977 and 1983[2,11] and by 1986 declined to 64.5/1000.[12]

After the revolution, Nicaragua was divided administratively into six regions and three "special zones." For health care administration, each is defined as a "health region" (see Figure 4). A health region is subdivided into many health areas, throughout which are distributed health posts, the first line of primary care. They refer patients needing more advanced care to a smaller number of health centers located

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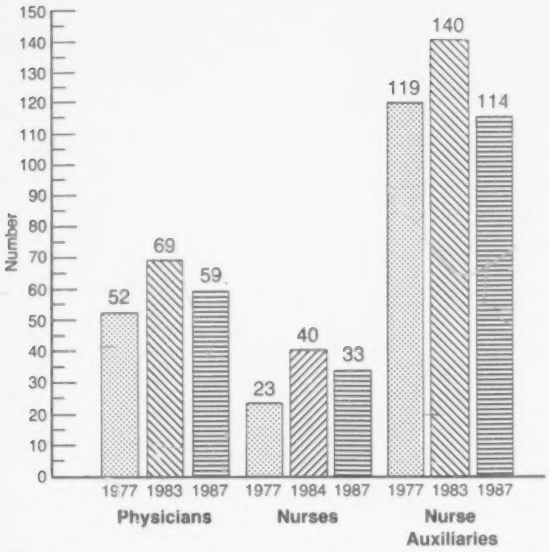

FIGURE 1—Increased Numbers of Physicians, Nurses, and Nurse Auxiliaries per 100,000 Population in Nicaragua since the Sandinista Revolution in 1979.
SOURCE: References 6, 8 and 10.

in urban areas. Patients requiring hospital care go first to municipal hospitals, located in the larger towns. These in turn refer to the regional hospital, located in the principal city of the region. Managua and León, which have the nation's two medical schools, also have the national referral hospitals that offer tertiary care.

A typical health post is managed by a nurse, has a doctor at least part-time, handles acute and chronic illnesses, pre-

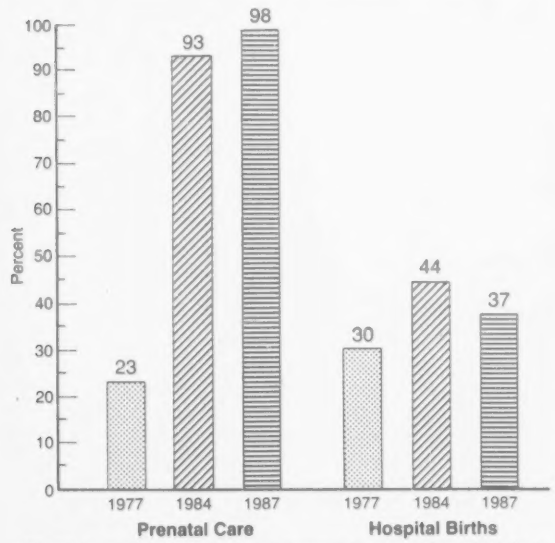

FIGURE 3—Increased Percentage of Prenatal Care and Hospital Births in Nicaragua since the Sandinista Revolution in 1979. (Includes care given by nurses and midwives.)
SOURCE: References 6, 8 and 10.

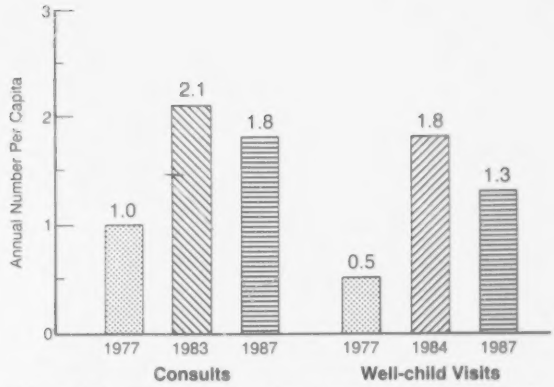

FIGURE 2—Increased Outpatient Consults and Well-Child Checkups for Children Under Age Five in Nicaragua Since the Sandinista Revolution in 1979.
SOURCE: References 6, 8 and 10.

natal checks, well-child visits, vaccinations, and minor suturing and wound care. Sometimes an Oral Rehydration Unit is included, where babies and children with mild to moderate degrees of dehydration are observed and given packets of salt and sugar containing liquids (the World Health Organization formula, provided by UNICEF). There is no clinical laboratory testing. No appointment is required. Special efforts have been made to place health posts in the most underserved rural and urban sectors.

Health centers are staffed by doctors who may be

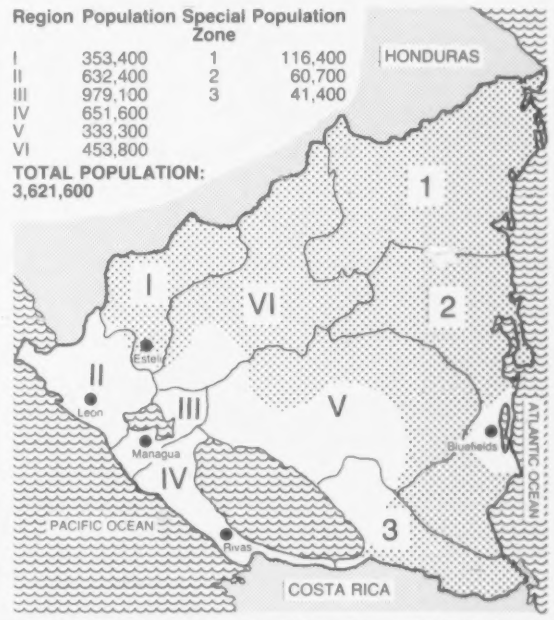

FIGURE 4—Map of Nicaragua Showing Division of Country into Six Regions and Three Special Zones, along with Current Population Data.
SOURCE: Reference 12.
NOTE: Each region or zone is defined as a "health region" and is further subdivided into "health areas" (not shown). Shaded area of map indicates war zones.

physicians just out of internship and completing two years of obligatory government service (an obligation common throughout Latin America) or by specialists (pediatrician, internist, etc.) completing residency training. There is on-site laboratory and x-ray equipment in a few health centers, and dentistry may also be offered. Some health posts and centers also have beds, where patients may be held for short periods of observation and treatment. In isolated areas, vaginal deliveries are performed in health centers. Patients are seen both by appointment and on a walk-in basis.

Municipal hospitals are staffed mainly by newly graduated generalist physicians completing their obligatory service, while regional hospitals have some specialists and, as teaching facilities, have intern staff as well. Although most of the physicians in outpatient facilities are Nicaraguan, there continues to be a shortage of Nicaraguan hospital physicians. Thus, in Region I (Estelí), more than 50 per cent of the hospital physicians in 1986 were international volunteers, mainly from Cuba but also from Western Europe, Canada, the United States, and other Latin American countries.* Of the 637 specialists in Nicaragua in 1987, 103 were foreign volunteers, nearly half from Cuba.** Cuba also provides generalist physicians completing obligatory postgraduate service who help staff health posts and centers as well as hospitals.

In practice, the system may not always function as planned. Although utilization of community-based facilities for primary care has increased greatly,[2] it is common in cities to see long lines outside hospital emergency rooms at 8:00 am while neighboring health posts may have only a trickle of patients. Some patients avoid the health posts and the centers and go directly to the hospital for ambulatory care because they feel the doctors may be more specialized (and thus better) or that more medicines or laboratory tests are available. While this is often true, the fact remains that, even in the hospitals, specialists are in short supply and medicines and laboratory tests are frequently unavailable.

People with money continue to utilize the private medical sector, which may provide up to 15 per cent of outpatient medical services.[3-5] In Nicaragua, the medical economy like the general economy is a mixed one; as in many developing countries, physicians tend to work mornings in government institutions and afternoons in their consultorios privados (private offices). Equipment and procedures absent in public medicine can often be found in private clinics, laboratories, pharmacies, and hospitals. In Estelí, private laboratories could often provide many clinical chemistry, blood cultures, and antibiotic sensitivity tests unavailable in the government hospital. Such services, though inexpensive by developed country incomes, cost at least the equivalent of several days' salary for an average Nicaraguan worker. Outside Managua and León, subspecialists such as urologists or otorhinolaryngologists are rarely found in the public sector, yet are readily accessible in private consultorios. Their fees limit access to those who can afford to pay them; sometimes payment in dollars is demanded.

For others, especially rural peasants and farm workers (campesinos) and people with less education, the traditional health care system of herbalists, curanderos, and lay midwives continues to be an alternative. Patients told me various reasons for consulting traditional healers: long-standing belief in their methods and medications, lack of access to more modern medical care in areas that are remote or destabilized by the contra war, and fear or distrust of doctors and hospitals.

*Sanchez E: Director of Planning, Region I Ministry of Health, Estelí, Nicaragua. Personal communication, 1986.

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Diagnosis and Treatment

Although tremendous efforts have been made to increase the number of doctors and nurses, the demand for medical care is overwhelming now that free health care is deemed a right for everyone. Beyond the initial step of increasing the quantity of health workers is the even more difficult challenge of improving their training and work environment. The lack of clinical teachers and attending physicians, already acute because of the exodus of well-trained doctors after the revolution, is worsened by the contra war, which has caused diversion of resources and manpower, including medical manpower.[2] Finally, the poverty of the country, exacerbated by the enormous costs of the war, has restricted all efforts to improve the quality as well as quantity of medical attention.

Outpatient Care

Physicians in outpatient clinics I observed in five of the six regions and one of the special zones often see large numbers of patients—up to 50 in a four-hour period. Foreign physicians working in different areas confirmed my impressions that physical examinations are often rudimentary and examining rooms are commonly dark, noisy, and shared by several practitioners, separated only by small partitions. Most patients spend only a few minutes with the doctor and then receive four or five prescriptions.[13] Charting is rarely done, except for minimal obstetric and well-child records. Medicines are virtually free, but quantities are limited.

Outpatient medical care is uneven and variable in its effectiveness. Patients get little information or health education from physicians, leading to frequent noncompliance or miscompliance. Other foreign physicians as well as I have witnessed the unnecessary or inappropriate prescription of drugs and the inadequate or inappropriate treatment of chronic conditions. Much of the impetus for this polypharmacy comes from the patients themselves. The beliefs, so prevalent in Latin America, that more medicine is better, and that injections and infusions are more powerful than pills, create expectations among patients that physicians find difficult to resist.

Nevertheless, the majority of every-day problems are adequately diagnosed and treated. The Oral Rehydration Units are one of the great successes of outpatient care, and many a child has been spared hospitalization because of this widely available service; more severely dehydrated children are detected and referred to hospital in a timely manner. Extensive popular education campaigns have played a key role in the development of this program.[13]

Hospital Care

Hospital conditions vary enormously. In Estelí, the hospital is over 40 years old and was partially destroyed by the National Guard in 1979. Water is shut off in most of the hospital during much of the day; lights sporadically go off during surgery and the back-up system sometimes fails. The wards are large dark chambers; occasionally two people share a bed. One the other hand, cities like Bluefields (on the Atlantic coast), and Matagalpa and Rivas have newly built modern hospitals with some semi-private rooms, up-to-date

sterilizing equipment, steam laundry, overhead paging, and new x-ray machines.

My own observations of many Nicaraguan hospitals and the reports of other foreign physicians working in these institutions indicate that, despite the limited resources, common problems are usually handled adequately if not optimally. Problems arise with more complex illnesses, however, because of lack of training or diagnostic capability. Patients of any age with unclear diagnoses tend to be treated vigorously with multiple antibiotics and frequently with steroids. Emergency suturing is often done by people with little training, even for complex cosmetic lacerations, so that the wounds are closed but the scars are unsightly. Crisis situations are often poorly managed. Many of the worst problems of mismanagement come from the smaller hospitals that are staffed only by newly graduated physicians with no back-up.

Regional and municipal hospitals have no facilities or personnel for psychiatric patients. When acutely psychotic patients are hospitalized, they are placed in the medical wards with other patients. Hospital workers and physicians, with little training in psychiatric illness, are often reluctant to deal with them. The outpatient-based mental health workers (psychologists and social workers) make daily rounds and medicate these patients, but control methods are poor, so that such patients are sometimes tied to their beds or roam freely about the premises. In contrast, the national psychiatric referral hospital in Managua has succeeded in lowering its formerly overwhelming patient census by opening day treatment centers around the city and helping families to care at night for patients who are family members. This program will be applied nationally as soon as resources permit.[14]

Numerous factors contribute to the occurrence of clinical disasters. Medical training is spotty, because of limited resources. Medical students often have access only to outdated texts; many of their clinical teachers are social service doctors just out of internship and assigned to the medical schools. Basic skills in history–taking and physical examinations are limited. Experience with laboratory procedures is minimal so that many physicians regard any laboratory result as gospel, no matter how inconsistent it is with the clinical picture. Moreover, laboratory results are often unreliable and such basic tests as blood typing, hematocrit, and pregnancy testing are often unobtainable because of lack of reagents or equipment. Only a few clinical chemistries are available; bacteriology is limited to staining, serology to VDRL. Pap smears and biopsies require two to four months turnaround time (only León and Managua have pathologists) and, even so, approximately 25–30 per cent of the requests are never returned.

Other components of hospital care contribute to suboptimal results. Nurses and nursing aides, with little training in physical diagnosis, sometimes fail to recognize the ward patient who is deteriorating or developing complications. Vital signs are frequently done casually, and accurate fluid balance records are very hard to obtain.

The referral hospitals in Managua and León have somewhat more resources than the regional facilities, but intensive care units lack functioning monitors, blood gases are usually unobtainable because reagents are lacking, and diagnostic imaging capabilities include only basic x-ray and "sometimes-available" ultrasound. Chemotherapy or radiotherapy for cancer is limited, and no oral narcotics are available for terminal patients. Some medical problems can be diagnosed and treated only in other countries. Although Cuba and some European countries accept such patients, the process of referral is complicated, bureaucratic, and out of the reach of most patients.

The problems, errors, and shortcomings that affect medical care in Nicaragua are serious. Misdiagnosis and mismanagement are a significant cause of current morbidity and even mortality. Nevertheless, there is no question that most people have benefited enormously by the free medical care that is now available to all. The occasional dramatic failures are countered by the common unspectacular successes.

Linkage with the Preventive Sector

The outpatient health facilities should be a key link between curative and preventive health services because they provide both types of service, especially in the area of maternal and child health. Child nutrition has received special focus, and the health posts and centers seek to identify malnourished children, educate the mothers in such cases, and enlist them in special programs to receive supplementary food grants. For sheer lack of resources and trained personnel, especially health educators and social workers, results are uneven and it is not uncommon to see the same malnourished child return many times to the clinic or hospital. Clinic nurses, auxiliaries, and physicians have not been prepared for the preventive and educational aspects of health care; moreover, they are overwhelmed by other tasks. As an example, while oral contraceptives and IUDs (intrauterine devices) are dispensed, women learn little about their bodies and physiology or about the contraceptive methods they are given. In other programs, such as prevention of cervical cancer (the leading cancer in Nicaraguan women), the delays and frequent loss of Pap smear results minimize effectiveness.

There is no link between hospitals and preventive care. Part of the problem is lack of a standardized referral system. Patients often leave the hospital with little understanding of what happened to them during their stay. They receive brief written summaries (often just the discharge diagnosis) and are told to get follow-up care in the health post or center. Minimal patient education is done by hospital personnel, and there are no educators working in the hospital comparable to those working in the community where specially trained volunteers ("brigadistas") have been an important strategy of the Ministry of Health.[3–5] Social services are understaffed, making it difficult to obtain home evaluations for children with frequent hospital admissions for diarrhea and malnutrition.

One positive area is the promotion of breastfeeding. This was one of the first health priorities after the revolution, and postpartum women are given all available aid and encouragement to help them establish lactation, even when their babies are temporarily unable to nurse. Rooming-in was established in maternity wards in the early post-revolution years and the promotional distribution and advertising of commercial infant formulas were banned. Little data are available to support the general impression that more women are breastfeeding, however. Results of the few existing studies suggest that many women revert to bottle-feeding within the first four months.[15,16]

Discussion and Conclusions

Nicaragua, in the nine years since the Sandinista revolution, has developed a medical system that is physically and

financially accessible, offers care that is very uneven in quality, is generally adequate for most common problems, and is suboptimally coordinated with preventive health efforts. Such conclusions, however, must be viewed in a larger framework. It is vital to compare the present situation with pre-revolutionary conditions and with health care in the rest of Latin America to appreciate the achievements as well as understand the problems.

The present system is an immense improvement for the great majority of Nicaraguans, who previously had limited or no access to health care. Quality of care has probably decreased for those with wealth, but even in the Somoza epoch the wealthy (including the Somozas themselves) were quick to seek medical care outside Nicaragua when faced with serious illness.

Other Latin American countries have far more sophisticated medical capabilities than Nicaragua but, except for Cuba and Costa Rica, there is no other country that has managed to offer free curative health services throughout the country that reach even the most rural areas. The usual pattern is that physicians cluster in the cities and that services are often limited by access or the ability to pay.

Nicaragua in the 1970s was among the poorest of Latin American nations. Since the revolution, up to 10.9 per cent of the national budget has gone for health care,[10] but the poverty of the country continues to limit efforts to upgrade medical care. Phone communication is extremely difficult or impossible, paper is scarce, transportation is overcrowded, funds for equipment and maintenance are limited, and the situation is exacerbated by war.

Beyond the material deficits are the cognitive and psychological consequences of underdevelopment. Illiteracy was reduced from over 50 per cent to less than 13 per cent soon after the revolution,[1] but many people have only rudimentary reading and writing skills. Fatalism and inertia based on past experience must be overcome; long-established but questionable hospital routines require change and patients with chronic disease must be convinced to continue therapy.

The American economic embargo imposed on Nicaragua in 1985 by President Ronald Reagan has had major effects on the health care system. Prior to the embargo, most medicine and medical equipment imported into Nicaragua was American. Although the embargo supposedly excludes medical material, in practice most American companies and shippers are no longer supplying Nicaragua. The consequences are readily visible: Kodak x-ray film processors stalled for lack of chemicals, blood-gas machines inactivated by lack of parts, American microscopes unusable without replacement bulbs. Moreover, the costs of the economic sanctions to the Nicaraguan economy have meant less funds available for health.

The effects of the contra war on the health care system include the intentional destruction of facilities and deliberate killing of health personnel.[2,11,17-19] The war has also diverted resources destined for health care, displaced large numbers of people, and disrupted health services in the northern areas of the country.[2,11,20] In Estelí, I attended women arriving from war zones with complications of high-risk pregnancies who had been unable to obtain ongoing prenatal care for their underlying diabetes, renal insufficiency, or hypertension because the health facilities in their area had been destroyed or travel was dangerous because of contra mining of roads.[21] Estelí itself had been scheduled for construction of a new hospital in 1986; the site had already been prepared but the costs of the war forced cancellation of the project.

Even without the war, improving curative health care in Nicaragua would have been more difficult than developing the public health sector. The minimal existence of a public health system in Somoza's time allowed the new government to build a new one almost from scratch, whereas it was necessary to contend with an already entrenched medical care system and hierarchy when planning and implementing new ways of delivering medical care. In addition, curative services require a much greater investment in training and technology than preventive services, so advances are necessarily slower.

Nicaraguan health officials, cognizant of all the above mentioned problems, continue to apply new approaches to better the system, especially to improve clinical training and continuity of care.[22] The internship year, previously spent entirely in the hospital, was recently changed to add a 10-week block of training in a health center because internship is followed for most graduates by two years obligatory service in outpatient facilities. Health centers in Managua are now being kept open 24 hours a day to relieve the burden on hospital emergency rooms. Standardized forms have been introduced to summarize the obstetric history and outcome of parturient women and to specify and record vaccination schedules for their infants; these forms are distributed to women after delivery. In Estelí, a government-sponsored Center for Natural Childbirth offers free prenatal classes preparing women and their partners with Lamaze and other instruction for a better birthing experience. It has been so successful that similar centers will be opened around the country. Also in Estelí, a "Rescue of Popular Medicine Project" has involved the systematic classification and evaluation of herbal medicines and other natural remedies to determine whether they are effective enough to replace imported pharmaceuticals, thus saving precious foreign exchange. This project also is being expanded to other regions. A carefully constructed National Medical Formulary of less than 500 agents was instituted in 1985 to foster more rational and economic distribution and utilization of conventional pharmaceuticals at different levels of the health care system.

Perhaps the most important development is a new residency training program in "medicina integral." Health planners were concerned that too many specialists were being trained and that three-fourths of all practicing physicians were hospital-based, with excessive concentration in the most urbanized regions (León and Managua).[23] In response, they developed a three-year program that borrows heavily from Western family practice concepts but also includes emphasis on public health and administrative training. The residency training will aim at producing physicians with broad skills who can manage the municipal and regional hospitals and the larger health centers as well as staff isolated rural facilities. These physicians will be trained in the management of common problems, including surgical procedures such as cesarean sections and appendectomies, but they will also be able to recognize, stabilize, and refer more complex conditions, including acute psychiatric illness. The curriculum begins with an intensive block designed to improve skills in history–taking, physical examination, laboratory utilization, and appropriate medication prescribing. The long-range plan is to make these new "médicos integrales" the cornerstone of the system, thereby improving the distribution and continuity of care.

On a personal level, doing medical work in Nicaragua is

deeply gratifying despite the limitations and frustrations. Socioeconomic status is not a consideration in treating patients, nor is there any subtle prejudice against the poor. There is remarkable patience in the face of constant inconvenience and hardship. A democracy of scarcity prevails: patients of all walks of life are together on the wards, while hospital personnel share all facilities equally, including meals in the same dining hall. Young physicians are eager for teaching and support. In general, there is a sense of unity and common purpose among health workers and patients that helps everyone tolerate the difficult conditions.

Improving curative health care in third world countries has received much less attention in recent years than improving public health. Still, the need for medical care will always exist, and the availability and adequacy of such care is highly visible to the general population, which may influence public confidence and participation in preventive efforts. New strategies are necessary for developing nations seeking to create, with limited resources, health care delivery systems truly responsive to national needs. Despite the obstacles of poverty and war, Nicaragua has made considerable progress in this area. For a poor person in need of medical care, Nicaragua today is a much better place than eight years ago and better for such a person than a great many other Latin American countries.

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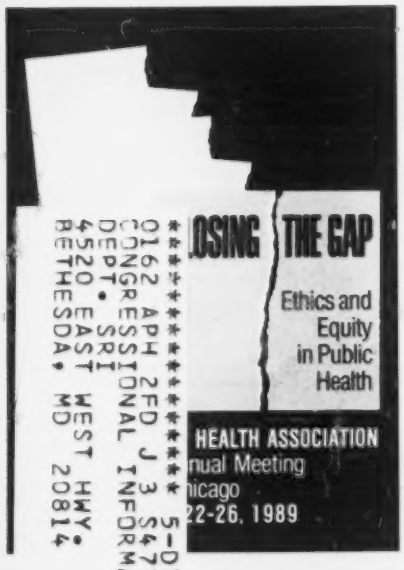

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Use of Clinic versus Private Family Planning Care by Low-Income Women: Access, Cost, and Patient Satisfaction

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Abstract: Use of private physicians versus public family planning facilities by poverty level and near poverty level women was examined by means of a sample survey conducted in low-income areas of Los Angeles County. Utilization differed by race/ethnicity, with Hispanics more likely to go to federally subsidized family planning clinics (primarily county-run), Whites and Blacks to private physicians. Private family planning offers easier access, greater convenience, and higher satisfaction, albeit at almost double the cost. Clinic usage is influenced by lack of a regular source of medical care and lack of insurance coverage more than poverty level per se. Clinic patients report greater patient education regarding contraceptive methods, but less general medical care during clinic visits. They are more likely than private patients to express a desire for a different source of family planning care. (*Am J Public Health* 1989; 79: 692–697.)

Introduction

A shift over the course of the last two decades toward the use of more effective types of contraception requiring a physician or clinic visit[1] has fostered increased attention to utilization of formal family planning services, particularly for low-income populations and other subgroups with higher than average fertility rates.[2] For approximately 15 years, federally subsidized family planning clinics have been providing care at low cost to the patient. These clinics provide contraception and other services primarily to low-income populations[3] that might otherwise have limited access to reproductive-related care.[4] Yet, data from cycle II of the National Survey of Family Growth (NSFG) show that the percentage of women below the poverty level who obtain family planning care from private sources (rather than from subsidized clinics) ranges from 40 percent and 44 percent for Hispanics and Blacks, respectively, to 75 percent for Whites.[5]

The use of private sources of family planning care by a substantial minority of Black and Hispanic women below the poverty level, along with the pronounced difference in use of private care by poverty-level Whites, suggests the presence of determinants of private versus public sources of care other than poverty level per se. Cycle III of the NSFG (which incorporated both married and unmarried women) reported that provider choice was influenced by insurance coverage, with women who had no insurance (private or public) more likely to use subsidized clinics.[6]

Factors predictive of sources of family planning care, however, do not fully explain differences in utilization. In this study, survey data stratified by poverty level (below poverty, 100–149 percent of poverty, and 150+ percent poverty) and race/ethnicity (White, Black, Hispanic) are used to examine demographic precursors of the choice of provider for family planning care. We examined factors related to access, cost, and satisfaction with care received from clinic and private sources, as well as reported levels of patient education provided by different sources of care and the comprehensiveness of medical care available at each source.

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Methods

Study Population

We conducted a population-based survey of poverty level and near-poverty level women residing in Los Angeles County. The target population was defined as women 18–44 years of age, or under age 18 if married or cohabitating. The survey employed a stratified random sample, with strata defined by race/ethnicity (White, Black, or Hispanic) and poverty level (below poverty, 100–149 percent poverty, or 150+ percent poverty). In order to ensure sufficient numbers of respondents from the two lowest poverty level strata, sample census tracts were drawn from among those in which more than 60 percent of the households were below 200 percent of poverty in the 1980 census.

Data Collection

A "screening questionnaire" determined participant eligibility and served as the basis for categorizing respondents into the sampling strata enumerated above. All respondents had visited a doctor or clinic for family planning care within a three-year period prior to the time of the interview. The screening questionnaire also served as the basis for stratifying the sample into equal numbers of White, Black, and Hispanic women, as well as equal representation of the three poverty level strata. The Asian population was excluded. Poverty level was determined by income/family size charts based on federal poverty criteria.[7]

Female interviewers conducted the interviews with respondents at assigned addresses in each census tract by means of a detailed survey instrument. The survey covered utilization of family planning services as well as other areas related to contraception and reproductive health. The screening questionnaire assisted the interviewer in identifying an eligible respondent (if any) within the household and selecting one woman in the household if more than one was eligible.

The average annual income for households below the poverty level was $7,348; for those from 100 to 150 percent of povery, $11,794; and for those in the open-ended 150+ percent poverty stratum, $22,309. Racial/ethnic differences in income occurred only in the third poverty stratum, where average income ranged from $18,322 for Blacks to $25,319 for Whites. When a respondent met all eligibility requirements and fell into a survey stratum not yet filled by the required number of respondents, the main interview was conducted. Field work for the survey was conducted between November 1984 and April 1985 and resulted in a total of 454 respondents—a response rate of 83.1 percent. The 16.9 percent nonresponse is comprised of refusals by eligible respondents (7.5 percent) combined with an

additional 9.4 percent "indirect" refusals representing non-screened households in which residents were inaccessible due to a security building, were not home after repeated attempts, or refused to provide screening information. (The latter figure was determined by multiplying the number of non-screened households by the probability of finding an eligible respondent in a screened household).

Survey Variables

The survey interviews, 29 percent of which were conducted in Spanish, cover the respondent's usual source of medical care, her insurance coverage and a number of other health and fertility-related areas. The "source" of family planning care was identified based on the respondent's report of where she went on her most recent visit for family planning care or advice. For approximately 80 percent of respondents, this visit had occurred within the preceding year.

From respondent information regarding name of facility, location, and the physician's name, sources of care were identified as one of four types.

- Subsidized clinics receive state and federal funding through the Los Angeles Regional Family Planning Council. Composed primarily of large, county-run clinics, but also including such organizations as Planned Parenthood, these clinics provide family planning care to low-income women on an income-based sliding scale.
- Private clinics receive no subsidy but typically accept Medicaid payment. These include both specialty and general clinics where patients do not have a regular physician.
- Health Maintenance Organizations (HMOs) are prepaid plans such as Kaiser Permanente, Maxicare, etc.
- Private physicians are non-prepaid doctors in solo or group practices.

Care facilities out-of-state or in Mexico were designated as "other."

A substantial portion of the interview elicited information on characteristics of the family planning care received at the respondent's source of care, including both objective and subjective measures of patient satisfaction, patient education, provision of preventive care, and cost. The analysis presented here contains comparisons of these utilization and satisfaction measures across clinic and private sources of family planning care.

Results

The distribution of respondent characteristics by race/ethnicity and poverty level (Table 1) reveals that: respondents in the highest income group are oldest; Hispanics are most likely to be married; high parity is associated with Hispanic ethnicity, and with incomes below the poverty level. Expected racial/ethnic differences in religious preference were also demonstrated.

Table 2 shows that 25.8 percent of the study's respondents are currently utilizing subsidized family planning clinics, 15.6 percent use private clinics, 18.5 percent use HMOs, 35.5 percent use physicians in private practice, and 4.6 percent use sources in other states or Mexico. This pattern of utilization does not vary by poverty level, but 41.9 percent of

TABLE 1—Respondent Characteristics by Race/Ethnicity and Poverty Level

Respondent Characteristics	Race/Ethnicity			Poverty Level			
	White	Black	Hispanic	Below Poverty	100%–149% Poverty	150%+ Poverty	Total
	%	%	%	%	%	%	%
Age (years)							
17–20	13.9	10.9	12.3	10.6	19.2	7.0	12.3
21–25	31.4	34.8	31.3	34.4	27.8	34.5	32.4
26–30	23.4	34.1	26.3	30.0	27.2	26.1	27.8
31–35	16.8	13.0	19.0	15.6	17.9	16.2	16.5
36–44	14.6	7.2	11.2	9.4	7.9	16.2	11.0
Total	100.0	100.0	100.0	100.0	100.0	100.0	100.0
Marital Status							
Married	36.5	22.5	62.0	29.4	43.7	54.9	42.3
Informally married/living together	19.7	13.0	20.1	16.3	16.6	21.1	17.8
Widowed	0.7	2.2	0.0	0.0	2.0	0.7	0.9
Divorced	8.0	14.5	1.7	9.4	7.9	4.9	7.5
Separated	10.2	8.0	7.3	15.6	5.3	3.5	8.4
Never married	24.8	39.9	8.9	29.4	24.5	14.8	23.1
Total	100.0	100.0	100.0	100.0	100.0	100.0	100.0
Parity							
Zero	35.8	23.2	7.3	7.5	19.9	36.6	20.7
One	23.4	25.4	30.7	23.8	29.1	28.2	26.9
Two	18.2	26.8	29.1	23.1	31.8	20.4	25.1
Three	13.9	13.0	16.2	21.3	11.9	9.9	14.5
Four or More	8.7	11.6	16.7	24.3	7.3	4.9	12.8
Total	100.0	100.0	100.0	100.0	100.0	100.0	100.0
Religious Preference							
Protestant	35.3	73.7	8.9	38.8	40.9	28.9	36.5
Catholic	22.1	13.9	85.5	46.3	47.0	40.8	44.7
Jewish	4.4	0.0	0.6	0.6	0.0	4.2	1.5
Other	3.7	2.2	0.0	1.3	2.0	2.1	1.8
None	34.6	10.2	5.0	13.1	10.1	23.9	15.5
Total	100.0	100.0	100.0	100.0	100.0	100.0	100.0

TABLE 2—Source of Family Planning Care by Race/Ethnicity and Poverty Level

Total	Source of Care					
	Subsidized Clinic (N = 117)	Private Clinic (N = 71)	HMO (N = 84)	Private Physician (N = 161)	Other (N = 21)	Total (N = 454)
	%	%	%	%	%	%
Total	25.8	15.6	18.5	35.5	4.6	100.0
Race/Ethnicity						
White	18.2	12.4	23.4	43.8	2.2	100.0
Black	12.3	21.0	23.2	39.1	4.3	100.0
Hispanic	41.9	14.0	11.2	26.3	6.7	100.0
Poverty Level						
Below Poverty	31.7	11.8	19.3	34.2	3.1	100.0
100%–149% Poverty	24.5	19.2	15.2	35.1	6.0	100.0
150%+ Poverty	20.4	16.2	21.1	37.3	4.9	100.0

Hispanics utilize subsidized clinics, compared to 18.2 percent of non-Hispanic Whites and 12.3 percent of Blacks. Poverty level and race/ethnicity interact (Figure 1). Hispanics below the poverty level are twice as likely to utilize subsidized clinics compared to those above poverty; for Hispanics below poverty, these clinics account for almost two-thirds of all family planning care.

In Table 3, the percentage respondents with a regular source of medical care ("one person or place in particular you usually go to when you are sick or want advice about your health") and the percentage with insurance coverage are compared by source of family planning care and by race/ethnicity and poverty categories. The percentage reporting a regular source of medical care is substantially lower for those using both subsidized and private clinics compared to HMOs and private physicians, for all strata. Women who utilize subsidized family planning clinics are substantially less likely to have either private or public insurance coverage compared to those who utilize other sources. Less than one-third of subsidized clinic patients have any coverage, compared to half of those utilizing private clinics, more than three-fourths of those with private physicians, and more than nine-tenths of those who utilize HMOs.

Comparisons across race/ethnicity and poverty level

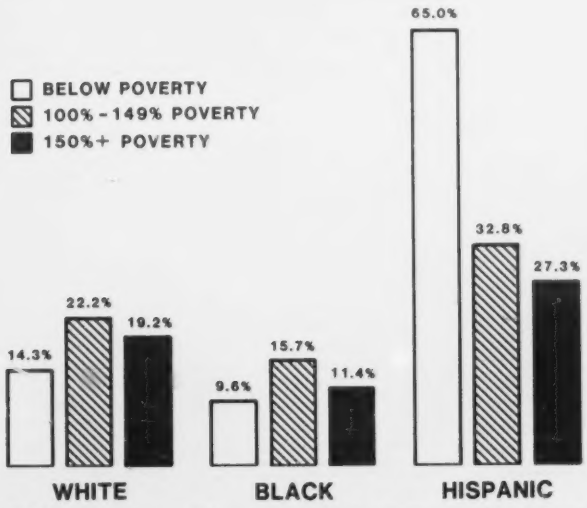

FIGURE 1—Percentage of Respondents within Racial/Ethnic and Poverty Level Categories Utilizing Subsidized Family Planning Clinics, Los Angeles County Survey

show differences in insurance coverage within each group between patients of private physicians and those who utilize clinics. Hispanics are substantially less likely to have insurance than are Whites and Blacks (47.6 percent compared to 72–73 percent). At 64 percent, the level of private/public health insurance reported by poverty level women is identical to that found for poverty level respondents to the National Health Interview Survey.[8] But the percentage of respondents with insurance does not increase significantly for those above the poverty level in this low-income population.

A number of other characteristics not shown in Table 3 were compared across a patient's source of family planning care and found to have a small or inconsistent relation to type of provider. These include the woman's age, her level of education, marital status, employment status, religion, length of residence in the Los Angeles area, and number of pregnancies, live births, and abortions.

Table 4 presents measures of access to family planning care, the physical examination during the most recent visit, and the cost for that visit, along with indirect measures of patient satisfaction. The percentage reporting that an appointment was required ranges from 95.2 percent for HMOs to 80.7 percent for private physicians, with private and subsidized clinics falling in between. The average number of days between the time the appointment is made and the date of the visit exceeds 12 days for both subsidized clinics and prepaid plans, compared to 5.1 days for private physicians.

An average of 156.9 minutes is reported as the total visit time for subsidized clinics, compared to 91.0 minutes for private clinics, and 68–70 minutes for HMOs and private physicians. Subsidized clinics also have higher reported numbers of people in their waiting rooms, and are less likely to permit women to bring their children, compared to other sources of family planning care. Greater difficulty reaching the woman's provider by telephone is encountered with subsidized clinics, followed by HMOs, whereas private physicians and clinics are easiest to reach by telephone.

A series of questions designed to measure respondents' satisfaction with their source of family planning care (e.g., "I am usually able to come in at a time and date that are convenient for me") show highest rates of agreement (either "agree" or "strongly agree") with most of the favorable statements by patients of private practitioners, and highest rates of agreement with unfavorable statements on waiting times and crowdedness by patients of subsidized clinics (Table 4). Statements on costs and return visits did not differentiate among providers of care.

The actual cost of family planning care (Table 4) is reported

TABLE 3—Percentage of Respondents Reporting a Regular Source of Medical Care, Insurance Coverage by Source of Family Planning Care, Race/Ethnicity and Poverty Level

	Source of Care				
	Total[a] (N = 454)	Subsidized Clinic (N = 117)	Private Clinic (N = 71)	HMO (N = 84)	Private Physician (N = 161)
Percentage with Regular Source of Medical Care	%	%	%	%	%
Overall	72.5	61.5	64.8	81.0	82.0
Race/Ethnicity					
White	63.5	40.0	52.9	68.8	75.0
Black	83.3	70.6	69.0	87.5	92.6
Hispanic	70.9	66.7	68.0	90.0	78.7
Poverty Level					
Below Poverty	68.3	58.8	47.4	67.7	87.3
100%–149% Poverty	73.5	62.2	69.0	91.3	81.1
150% + Poverty	76.1	65.6	73.9	86.7	77.4
Percentage with Health Insurance					
Private insurance/health plan	41.2	21.1	36.6	67.9	43.5
Medicaid or other public funds	22.1	9.6	15.5	23.8	32.9
Total: private and/or public insurance[b]	62.8	30.7	49.3	91.7	76.4
Race/Ethnicity: Total Insurance					
White	73.1	60.0	52.9	96.9	71.7
Black	71.5	26.7	48.3	87.5	87.0
Hispanic	47.6	21.6	48.0	90.0	70.2
Poverty level: total insurance					
Below poverty	63.6	26.5	52.6	90.3	85.5
100%–149% poverty	56.7	38.9	37.9	87.0	66.0
150% + poverty	68.1	27.6	60.9	96.7	77.4

[a]Includes 21 respondents who reported "other" sources of care.
[b]These percentages are slightly lower than the sum of "private" and "public" percentages, as a small number of respondents reported having both types of coverages.

as the percentage of patients with no "out-of-pocket" costs, and the average dollar amount for patients who incurred some direct cost for the visit. Of patients of subsidized clinics, 28.3 percent had no direct cost, compared to 55.1 percent of patients of private physicians. For those who had some out-of-pocket expense, however, the difference in that expense is almost double—$18.20 for subsidized clinics compared to $32.80 for private physicians. This comparison excludes those who had procedures (including sterilizations) performed during the visit, along with a small number of respondents who were being seen primarily for a medical complaint.

Data on characteristics of the physical examination (Table 4) show a four-fold difference between patients of subsidized clinics and patients of private physicians in the percentage who reported that they have the same examiner at each visit. Waiting times for the examination range from just over 30 minutes for prepaid plans and private physicians to over 90 minutes for subsidized clinics; however, the average length of the examination does not differ by source of care.

Patient satisfaction items related to the physical examination show higher agreement with favorable communication/discussion statements by private practice patients, and higher levels of discomfort and embarrassment for subsidized clinic patients. Statements on completeness of the examination and consideration and gentleness of the examiner do not differentiate among providers. Finally, when respondents were asked if they would prefer a different source of care, 33.3 percent of subsidized clinic patients responded affirmatively, compared to 26.2 percent of those in HMOs and 15.5 percent of patients of both private physicians and private clinics.

Table 5 presents results from a question on the comprehensiveness of patient education on contraceptive methods, along with information on the source(s) of contraceptive patient education at the woman's source of care. Respondents were shown the list of family planning methods in Table 5 and asked to indicate whether each method listed had been discussed or explained to them at their source of family planning care. This family planning education could have occurred prior to the most recent visit, as virtually all respondents (91.5 percent) reported multiple visits within a three-year period to the place cited—a figure that did not differ by source of care. For each family planning method, subsidized clinics (followed in most cases by private clinics) are more likely to have provided an explanation than are HMOs and private physicians. Among the latter groups, approximately one-third of patients report having "no methods explained," compared to approximately 17 percent for subsidized or private clinics. Percentages for individual methods are likely affected by respondent recall, as subsidized clinics are required to provide information on all family planning methods. For those respondents who did report family methods explained during their visit, Table 5 shows the means by which the methods were discussed or explained. While the respondents' perceptions of types of health personnel may not be entirely reliable, it is clear that for those using HMOs or private physicians, the physician predominates as the source of information; whereas for subsidized and private clinics the source of information is more likely to be a nurse, nurse practitioner, counselor, or a film or videotape.

Finally, respondents were also asked if they thought that they could receive care for each of a list of common problems while obtaining family planning care. The list of problems was derived from data on the most common patient complaints in ambulatory primary care from the National Ambulatory Medical Care Survey.[9] Some complaints that suggested the possibility of a side effect of a birth control method were eliminated. Results of this comparison, not shown in the

TABLE 4—Access to Care, Characteristics of Examination, Satisfaction and Cost, by Source of Family Planning Care

	Source of Care			
	Subsidized Clinic (N = 117)	Private Clinic (N = 71)	HMO (N = 84)	Private Physician (N = 161)
Access to Care				
Percentage with Visits by Appointment[a]	90.6	84.5	95.2	80.7
Average Waiting Time for Appointment, in Days	12.4	9.9	12.9	5.1
Average Travel Time to Care, in Minutes	26.7	26.6	25.3	21.0
Average Total Time Spent at Source of Care, in Minutes	156.9	91.0	70.4	68.0
Average Number of People in Waiting Room	20.8	12.8	11.6	7.9
Percentage Allowing Women to Bring Their Children	71.8	90.1	96.4	96.9
Percentage Responses on Ease of Reaching Provider by Telephone				
Always easy	56.3	82.9	63.4	85.5
Sometimes difficult	24.1	8.6	32.9	8.2
Very difficult	12.5	7.1	1.2	3.8
Almost impossible	7.1	1.4	2.4	2.5
TOTAL	100.0	100.0	100.6	100.0
Satisfaction with Source of Care (percentage "agree" or "strongly agree")[b]				
I am usually able to come in at a time and date that are convenient for me	37.6	64.3	73.8	85.1
I often have to wait for a long time to be seen after I arrive there	71.8	50.7	38.1	27.3
The people who work there are courteous and helpful	83.4	94.2	89.1	96.9
The waiting rooms are often too crowded	72.7	47.2	46.5	23.6
I feel comfortable being around the other patients who are there	69.3	85.7	89.2	88.8
The people who work there try to understand how I feel	70.2	91.1	83.1	95.6
The family planning care I receive there is the best that is available	71.5	82.3	74.0	87.3
The cost of family planning care there is too high	15.8	10.5	7.7	16.3
The people who work there did not spend enough time explaining things to me	40.2	42.9	39.1	24.4
You have to go back too often for the same reason	16.4	15.9	10.8	9.5
Cost of Family Planning Care[c]				
Percentage with no Out-of-Pocket Cost	28.3	44.4	78.1	55.1
Dollars for those with Out-of-Pocket cost[d]	18.20	22.06	11.50	32.80
Characteristics of Physical Examination				
Percentage with same examiner every visit	22.2	19.7	44.0	88.8
Average waiting time to see examiner, in minutes	92.1	56.9	32.9	32.2
Average length of examination, in minutes	13.4	14.6	14.8	15.1
Satisfaction with Physical Examination (percentage "agree" or "strongly agree")[b]				
The examiner told me everything he or she was going to do before doing it	65.5	87.1	79.5	86.3
I felt uncomfortable while waiting for the examination	56.9	52.1	35.8	28.6
The examiner spent enough time discussing things with me	56.9	77.4	75.0	84.4
The examiner asked questions that were too personal	13.8	12.7	8.3	8.7
The examination was very complete	88.6	95.7	87.8	93.7
The examination was embarrassing	46.1	36.6	22.9	26.1
The examiner was kind and considerate	92.9	97.2	91.7	95.1
The examination was done gently	92.3	97.1	95.1	96.9
I felt free to ask the examiner all of the questions that I wanted to ask	76.5	81.6	84.6	90.7
Percentage Who Would Prefer A Different Source of Care	33.3	15.5	26.2	15.5

[a]Respondent usually has to have an appointment, as opposed to "walk-in" care.
[b]Possible responses to each statement were "strongly agree," "agree," "disagree" or "strongly disagree."
[c]Cost measures exclude patients who reported sterilization as their family planning method.
[d]Not including what insurance pays for. Cost comparisons exclude out-of-pocket costs reported by respondents who had any type of procedure (including sterilization) performed and those whose visit was primarily for some medical complaint (rather than primarily for family planning).

Table, produced consistent differences, with patients of HMOs most likely to believe that they could receive care or advice for each of the problems listed, followed by patients of private physicians, and (lastly) those who utilize subsidized or private clinics. Data on Pap smear, breast examinations, routine laboratory tests and tests for sexually transmitted diseases (STD) (also not shown) indicate no significant differences by source of care in the percentage of women who reported having had these screening examinations and tests performed within the preceding 12 months.

Discussion

Our data indicate that the majority of low-income women in Los Angeles County obtain their family planning care from private sources—primarily private physicians. Subsidized family planning clinics account for the majority of care only for the sub-group of Hispanics below the poverty level. Clinic utilization is associated with a lack of a regular source of general medical care, but the single most powerful predictor of patterns of utilization was insurance coverage, including Medicaid. Further, the relation of insurance coverage to race/ethnicity helps to explain high rates of subsidized clinic use by Hispanic women.

The study has thus demonstrated variation in the use of clinic versus private family planning sources by race/ethnicity, an absence of effects based on poverty level, and the impact of insurance coverage as a barrier to unconstrained choice of providers. These patterns of utilization by race/ethnicity and poverty level in Los Angeles differ from those from cycle III of the National Survey of Family Growth, which showed substantially greater utilization of subsidized clinics by Blacks and women below the poverty level.[6]

Both surveys found insurance coverage (either private or Medicaid, eligibility for which is dependent upon both income

TABLE 5—Family Planning Methods Explained by Source of Care

	Source of Care			
	Subsidized Clinic (N = 117)	Private Clinic (N = 71)	HMO (N = 84)	Private Physician (N = 161)
Methods Explained	%	%	%	%
Sterilization	44.4	46.5	33.3	40.4
Birth control pills	76.1	78.9	57.1	59.0
IUD	75.2	78.9	52.4	54.0
Diaphragm	70.1	71.8	51.2	47.8
Cervical cap	46.2	36.6	25.0	32.9
Condom or rubber	65.8	54.9	36.9	35.4
Foam, cream, jelly	65.8	40.8	32.1	36.0
Foam and condoms	60.7	53.5	35.7	36.6
Suppositories, vaginal inserts	47.8	31.0	19.0	26.1
Rhythm, natural family planning	45.3	33.8	27.4	28.0
Contraceptive sponge	39.3	22.5	15.5	24.2
"Morning after" pill	35.0	19.7	14.3	21.7
Withdrawal or "pulling out"	30.8	18.3	10.7	20.5
Douche or "washing out"	36.8	22.5	13.1	23.6
No methods explained	17.1	16.9	33.3	32.3
Methods Explained By				
Doctor	30.9	40.7	73.2	85.3
Nurse	44.3	35.6	19.6	11.9
Nurse Practitioner	8.2	15.3	3.6	2.8
Physician Assistant/Medex	1.0	3.4	1.8	0.9
Other Counselor	13.4	16.9	1.8	0.9
Film/Videotape	19.6	20.3	7.1	7.3
Pamphlet/Brochure	2.1	5.1	5.4	9.2

and legal residency status) to be a pivotal variable in explaining utilization of subsidized clinics versus private physicians. For women without insurance, subsidized clinics assume a crucial role in assuring access to family planning care. Differences in levels of utilization of subsidized clinics versus private care noted for the Los Angeles sample as compared to the national survey undoubtedly reflect regional variation in eligibility for Medicaid benefits used for private family planning care[10] as well as variation in both the organization and deployment of subsidized clinics, and the distribution of private practitioners across urban and rural communities.

The inclusion of a number of specific questions on convenience, cost, and satisfaction with the respondent's source of family planning care in the Los Angeles survey uncovered possible determinants of provider choice not considered in the national survey. Data on out-of-pocket expenses show that, although the cost of service from private physicians was almost twice that of subsidized clinics, over half of the patients of private physicians pay no out-of-pocket costs for visits as opposed to only about one-fourth of subsidized clinic patients, corroborating the role of lack of insurance coverage as a "barrier" to private family planning care.

Although there is no direct indication in our data of any differences in the quality of care among categories of providers, clinics are more likely to provide patient education for contraception. More importantly, a variety of patient satisfaction measures demonstrate why private family planning care will usually be sought by low-income women in Los Angeles when either private or public insurance is available. The detailed data on waiting times and other comfort and convenience factors associated with use of subsidized family planning clinics corroborate impressions obtained in a multicity telephone survey of low-income women, which showed that clinics were associated with less personalized and (according to respondent's criteria) "lower quality" care.[11] Our findings, although obtained from respondents in a single urban area, reinforce the importance of non-financial factors in determining levels of clinic utilization by low-income women, and point to specific areas (e.g., patient scheduling) in which efforts directed toward improvements in the provision of family planning services could be expected to yield benefits in terms of utilization and satisfaction.

ACKNOWLEDGMENTS

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Longitudinal Study of Physical Ability in the Oldest-Old

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Abstract: Based on 1984 data from the Longitudinal Study on Aging, one-third of White persons aged 80 or older living in the community (N = 1,791) were defined as having no difficulty in walking 1/4 of a mile, in lifting 10 pounds, in climbing 10 steps without resting, or in stooping, crouching or kneeling. Physical ability was associated with lower risk of death over two years mean follow-up; Relative odds (RO) = .4 (95 percent confidence interval = .4, .6) and in survivors, lower utilization of hospitals RO = .4 (CI = .3, .7), physicians RO = .6 (CI = .5, .8) and nursing homes RO = .3 (CI = .2, .5) compared with those having difficulty on any of the four functional measures included in the definition of physical ability.

Fifty percent of the women and 42 percent of the men physically able at the time of the baseline survey in 1984 remained physically able at follow-up. Continued physical ability in this group was associated with never having had cardiovascular disease RO = 2.1, (CI = 1.2, 3.7), never having had arthritic complaints RO = 1.9 (CI = 1.2, 2.7), a body mass index less than the 75th percentile (RO = 1.8 (CI = 1.2, 2.9), younger age (for each decade of age, RO = 2.0 (CI = 1.1, 3.6), and higher level of education (greater than 13 years versus 0–6 years) RO = 2.4 (CI = 1.2, 4.7). These correlates include factors amenable to preventive measures and highlight the need to consider the heterogeneity of the oldest-old in formulating programs aimed at prevention and postponement of disability. (*Am J Public Health* 1989; 79:698–702.)

Introduction

Increased attention is being focused on the fastest growing segment of the United States population, the oldest old.[1] Much of this attention concerns disability,[2,3] use of health services,[4,5] and risks of institutionalization.[6,7] Recently, more optimistic views of old age have been proposed by those focusing on "successful aging"[5] who suggest that environmental and physiologic modifiers of biological aging can be identified which may be amenable to preventive strategies even among the oldest-old. However, few studies exist describing the correlates of "successful aging" or the factors which may govern transition from health to disability for the oldest-old.

Data from a longitudinal follow-up of those persons age 80 or older participating in the Supplement on Aging to the 1984 National Health Interview Survey conducted by the National Center for Health Statistics were analyzed to address three areas pertinent to public health policy for the oldest-old:

1. What proportion of persons aged 80 or older in the community could be considered physically able and what is the change in the proportion physically able over time?
2. What is the relation of physical ability at baseline to morbidity and mortality over time?
3. Can we identify correlates of maintained physical ability which would be amenable to health promotion in the oldest-old?

Methods

Study Population

The National Health Interview Survey (NHIS) is a continuing survey of civilian noninstitutionalized persons in the United States which includes information on health status, health care utilization, health conditions, and disability. In 1984, the Supplement on Aging, specifically directed toward persons age 55 or older,[9] was added to NHIS; 96 percent of eligible persons participated. Special efforts were made to achieve complete self-response from all supplement respondents; of those for whom data were obtained, 89.8 percent were self-respondents. Permission was obtained to link with administrative records maintained by the National Center for Health Statistics. In 1986, a follow-up was undertaken of all persons aged 80 or older and approximately half of those aged 70–79 who participated in the Supplement on Aging. Vital status was ascertained for potential respondents.[10] Those alive at follow-up were reinterviewed over the phone with questions excerpted from the original survey instrument. Again efforts were made to obtain self responses; 74.3 percent of interviews were self-responses.

Our study population is drawn from the 2,090 persons aged 80 or older in 1984 who participated in the Supplement on Aging. Blacks and races other than Whites were excluded as numbers were few. Of the original 1,932 White persons, 1,776 were either reinterviewed or known to be dead as of the 1986 follow-up. Fifteen persons who were not interviewed at follow-up but who were known to be alive in institutions were included. This resulted in a final study population of 1,791 White persons aged 80 or older: 1,214 women and 577 men.

Study Outcomes

Physical ability at both baseline and follow-up was defined as no difficulty in any of four measures which have been associated with assessment of ability to work.[11] These four measures included having no difficulty in walking 1/4 of a mile, in stooping, crouching or kneeling, in lifting 10 pounds, or in walking up 10 steps without resting. Only individuals with no difficulty on all four of these measures were considered physically able. Individuals missing data on any of the four functional measures were grouped with the comparison group as were those persons not interviewed but known to be living in nursing homes (missing values ranged from 1.6 percent for walking to 7 percent for lifting in the baseline interview and from 3 percent for stooping to 5 percent for climbing steps in the follow-up interview).

Mortality at follow-up was defined as death validated by a match with the National Death Index[12] for deaths through August 31, 1986 or report of death for 13 people for whom

From the National Center for Health Statistics, Hyattsville, Maryland (Harris, Kovar, Kleinman, Feldman) and the National Institute on Aging, Bethesda, Maryland (Suzman). Address reprint requests to Tamara Harris, MD, MS, Office of Analysis and Epidemiology, National Center for Health Statistics, 3700 East-West Highway, Hyattsville, MD 20782. This paper, submitted to the Journal July 25, 1988, was revised and accepted for publication December 19, 1988.

Editor's Note: See also related editorial p 685 this issue.

information for a match was lacking. Death certificate information was not available. There were 338 deaths.

Three morbidity outcomes were defined for those alive and reinterviewed at follow-up. Hospital utilization was defined as two or more hospitalizations in the 12 months prior to reinterview in 1986. About 10 percent of the survivors had two or more hospitalizations; 5 percent were missing information on this variable. Nursing home utilization was defined as any nursing home utilization between the baseline interview in 1984 and reinterview in 1986. About 10 percent of the survivors had been or were presently in nursing homes and 3 percent were missing information on this variable. Heavy physician utilization was defined as six or more visits to the physician in the 12 months prior to reinterview (75th percentile). Twenty-five percent of the survivors reported six or more visits; 10 percent had missing data. Analyses presented in this paper exclude those with missing information on these morbidity outcomes.

To examine morbidity and mortality end points in relation to functional status, participants were classified into five mutually exclusive categories based on baseline functional status. The physically able group, the highest level of function, was used as the group against which risks were calculated. The four other categories included: those who received help from another person with one or more Activities of Daily Living (ADL) including bathing or showering, dressing, eating, getting in or out of bed or chair, walking, getting outside, using or getting to a toilet; those who had difficulty with Activities of Daily Living but did not receive help from another person; those who had no difficulty with Activities of Daily Living but did have difficulty with Instrumental Activities of Daily Living (IADL), i.e., preparing our meals, shopping for personal items, managing money, using the telephone, doing heavy housework, doing light housework; and those who were not physically able but had no difficulty on Activities of Daily Living or Instrumental Activities of Daily Living.

Change in Function

Those reporting a change in function for health or physical problem for difficulty walking a quarter of a mile or difficulty climbing 10 steps at follow-up were asked to define the main cause of change in function. These causes were compared for those physically able and the comparison group at baseline.

Correlates of Physical Ability

Variables examined in longitudinal relation to continued physical ability among those physically able at baseline included age, sex, social function (defined as phone contact with friends or neighbors within two weeks of interview), absence of hearing problems (defined as being deaf in one or both ears or having trouble hearing), absence of vision problems (defined as being blind in one or both eyes or having trouble seeing), no difficulty chewing, never having had hypertension, never having had diabetes, never having had arthritis or rheumatism, and never having had cardiovascular disease (stroke, heart attack, angina or other heart condition). Those responding "don't know" for any particular condition were considered to be free of the condition (less than 2 percent for any particular condition). Completed years of education were included as a categorical variable of 0–6 years, 7–8 years, 9–12 years and 13 years or more with 0–6 years as the comparison group. Missing values (2.5 percent) were imputed to the mean (8 years of education). Body mass index (weight in kilograms/height in meters squared) at baseline was calculated from reported weight and reported height and was included in the model as sex-specific quartiles of body mass index with the heaviest quartile as the comparison group. Missing values (less than 1 percent) were imputed to the mean for each sex, (23.7 kg/m^2 for men and 23.6 kg/m^2 for women).

Statistical Methods

The design effects, which represent the effect on the variance of using a clustered rather than a simple random sample, were small for each sex in this sample[9]; the inclusion of age in the models reduced the design effects further. The analyses in this paper are not adjusted for the complex sample design.

Logistic regression models were used to derive estimates of effect for analyses related to physical ability or to morbidity outcomes in those surviving from baseline to follow-up. All analyses were adjusted for age (as a continuous variable) and sex. Potential confounding by length of follow-up was controlled for by inclusion of a time variable calculated by subtracting the date of initial interview from date of follow-up interview. For analyses related to death, proportional hazards models were calculated with follow-up time defined as time from baseline interview to time of follow-up interview or date of death from National Death Index match. Mean follow-up time was 1.9 years. Survival analyses were not performed for risk factors for death among only those physically able at baseline due to the small number of deaths. Estimates of risk were derived from the coefficients in logistic or proportional hazards models and 95 percent confidence intervals were calculated.[13]

The independent variables were designed to express the association of a positive outcome with the unexposed state of a risk factor since the outcome in these analyses was physical ability—a positive outcome. For example, the relationship of physical ability to hypertension was assessed by contrasting risk in never hypertensives against risk in hypertensives, contrary to the customary expression of relative risk in epidemiologic studies.

Interactions with sex were observed for hearing problems ($p = .03$) and the social function ($p = .08$) variable. Final models included sex-specific indicator terms for these variables. No other significant interactions by sex were noted (P values ranged from .23 to .94); therefore, data are presented for men and women combined.

Results

Physical Ability at Baseline

In 1984, 33 percent of 1,791 noninstitutionalized White persons aged 80 or older, 44 percent of the men and 28 percent of the women, met the definition of physical ability (Figure 1). The order of difficulty for the four items included in the definition of physical ability was similar for men and women, although women were more likely to have difficulty on each item. Sixty-seven percent of all White persons aged 80 or older living in the community had no difficulty lifting 10 pounds, 57 percent had no difficulty climbing up 10 steps, 49 percent had no difficulty walking 1/4 of a mile, and 47 percent had no difficulty with stooping, crouching or kneeling.

Seventy-six percent of the physically able had had no hospitalizations and fewer than six doctor visits in the 12 months prior to baseline interview versus 55 percent of the comparison group, reflecting the better overall health status of the physically able persons. About 2 percent of the physically able had two or more hospitalizations in the 12

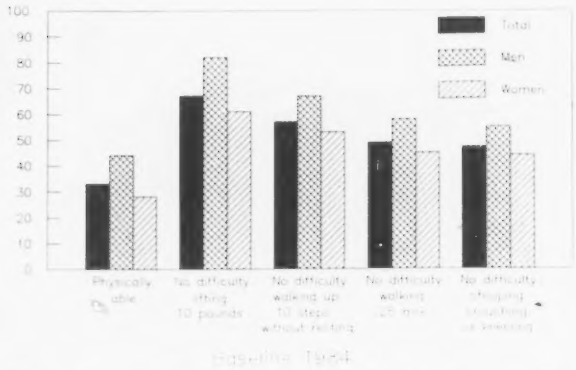

FIGURE 1—Physical Ability of Noninstitutionalized White Persons, Age 80 or Older

months prior to the baseline interview contrasted with 10 percent in the comparison group.

Longitudinal Follow-up

Fifty percent of the women and forty-two percent of the men who were physically able at baseline continued to meet the criteria for physical ability at follow-up (Table 1). The proportion of the comparison group dead by follow-up was twice that of the physically able group. Of the 1,453 persons alive at follow-up, 25 percent were physically able, 33 percent of the men and 22 percent of the women.

Cardiovascular conditions were the main problem reported to cause change in ability to walk 1/4 of a mile or in the ability to climb 10 steps among those physically able in 1984 (Table 2). Cardiovascular problems and multiple conditions were about equally important in terms of accounting for change in the comparison group.

Those receiving help with any ADL at baseline interview were four times more likely to die. Among the survivors, they were six times more likely to have used a nursing home during the following two years and at least twice as likely to have two or more hospitalizations or to have had six or more physicians visits in the year prior to reinterview than those physically able in 1984 (Table 3). Although risks of each outcome decreased as level of functional ability increased, even those only nonrobust without ADL or IADL difficulty were at increased risk in comparison with the physically able particularly for hospitalizations or nursing home admissions. When all those with missing data were considered as low

TABLE 1—Change in Physical Ability from Baseline to Follow-up

		Status at Follow-up		
Status at Baseline	n	Physically Able (percent)	Other (percent)	Dead (percent)
Men				
Physically able	251	42	43	15
Comparison	326	11	57	32
Women				
Physically able	343	50	43	7
Comparison	871	7	74	19
Total	1,791	368	1,085	338

TABLE 2—Main Problem Causing Change in Function

	Change in Walking 1/4 Mile		Change in Climbing 10 Steps	
Problems	Physically Able N = 84	Comparison N = 331	Physically Able N = 42	Comparison N = 263
	(Percent)			
Cardiovascular†	26	17	31	19
Multiple conditions	14	19	14*	21
Old age	14	14	14*	13
Arthritis	12	14	5*	14
Fall or hip fracture	14	11	14*	12
Other	20	25	22	21

*Less than 10 respondents in cell.
†Includes reports of circulatory disease, heart condition, hypertension, or stroke.

users of services, outcome results were substantially the same.

Correlates of Continued Physical Ability

Never having had cardiovascular disease and never having had arthritis at the baseline were associated with almost twice the likelihood of continued physical ability as was a body mass index below the 75th percentile (25.4 kg/m^2 for men and 25.2 kg/m^2 for women) (Table 4). Men were more likely to remain physically able. The same multivariate models showed higher education and younger age also were associated with continued physical ability. Several other characteristics were consistent with continued physical ability (no hearing problems in women, speaking on the phone with friends or neighbors in the two weeks prior to interview for women, no difficulty chewing, and no hypertension) although the confidence intervals for these variables did not exclude 1.0. When these analyses were recalculated considering those physically able individuals lost to follow-up (n = 41) as either physically able at follow-up or not, results were unchanged.

Discussion

Based on data from a national survey of noninstitutionalized older persons, about one-third of White persons age 80 or older living in the community had a high level of functional capability. This functional capability, which we call physical ability, appears to be associated with good health and low risk of morbidity and mortality over short-term follow-up.

Continued physical ability among those physically able at baseline was associated with lack of cardiovascular disease or arthritis, and moderate body weight. The prominence of cardiovascular disease and arthritis in these analyses was similar to data from other surveys[14,15] as well as data on bed disability, hospital utilization, and physician contacts.[16] Little data exist on the relationship between weight and morbidity for the oldest-old. It would be useful to evaluate our findings in a data set with measured weight and height. Our ability to explore other issues related to transitions in health status will be enhanced when linkages of the Longitudinal Study of Aging and Medicare records are in place and cause-specific mortality data are available from death certificates.

Although others have studied good health or longevity as an outcome in old age,[17,18] there have been few studies of longitudinal correlates of high level function among the oldest-old. A recent paper from the Framingham Heart Study

TABLE 3—Morbidity and Mortality Risks by Level of Function†

Status at Baseline	Death N = 338	Among Survivors: Any Nursing Home Admission N = 152	Among Survivors: Two or more Hospitalizations in Prior Year N = 150	Among Survivors: Six or more Physician Visits in Prior Year N = 368
Receive help ADL	4.4 (3.2, 6.1)	6.7 (3.8, 12.8)	3.3 (1.9, 5.5)	2.3 (1.6, 3.4)
Difficulty ADL—no help	1.9 (1.3, 2.7)	3.7 (2.0, 7.4)	2.0 (1.1, 3.4)	1.3 (0.9, 1.9)
Any difficulty IADL	2.2 (1.4, 3.3)	2.8 (1.4, 5.6)	2.5 (1.4, 4.5)	1.6 (1.0, 2.3)
Not physically able—no ADL or IADL	1.5 (1.0, 2.2)	2.2 (1.1, 4.1)	2.1 (1.2, 3.5)	1.3 (0.9, 1.9)
Physically able—no ADL or IADL	-1.0-	-1.0-	-1.0-	1.0

†Adjusted for age and sex (95 percent confidence intervals).

TABLE 4—Correlates of Continued Physical Ability at Follow-up

Variables	Relative Odds (95% confidence intervals) N = 513
Age 80 vs age 90	2.0 (1.1, 3.6)
Male vs female	3.2 (1.2, 8.6)
Education (years completed)	
0–6 years	1.0
7–8 years	2.2 (1.2, 4.0)
9–12 years	2.1 (1.2, 3.8)
13+ years	2.4 (1.2, 4.7)
Social contact—men	0.7 (0.4, 1.4)
Social contact—women	1.7 (0.8, 3.6)
No vision problems	1.1 (0.6, 2.1)
No hearing problems—men	0.6 (0.4, 1.2)
No hearing problems—women	1.5 (0.9, 2.5)
No difficulty chewing	1.3 (0.8, 2.4)
Never had hypertension	1.5 (1.0, 2.1)
Never had cardiovascular disease	2.1 (1.2, 3.7)
Never had arthritis	1.9 (1.2, 2.7)
Body mass index	
less than 25th percentile	1.8 (1.0, 3.2)
25–49th percentile	1.7 (1.0, 2.9)
50–74th percentile	2.1 (1.2, 3.6)
greater than 75th percentile	1.0

examined good function as an outcome over 21 years of follow-up in a group of 1,474 persons aged 35–68 at baseline.[19] Level of education was the only consistent predictor of good function for both men and women. That study, like ours, used physical function to define the outcome. Correlates of good function were examined in a group presumed to be at a high level of function at baseline. However, the Framingham study contained few very old persons at follow-up (n = 147 for those age 76 or older). Furthermore, their analysis defined long-term risks of good function in old age; our study examined short-term risks of remaining functionally able in very old age. As the US population becomes older, an understanding of how to remain healthy from younger life is critical; however, just as critical is an understanding of remaining as functionally fit as possible in the last years of life.

The consistent finding that level of education is related to good function is of interest despite the differences in study design and our inability to control for behavioral variables (such as cigarette smoking or alcohol consumption) or measured biological variables (such as blood pressure). Education may reflect life-long access to medical care or healthful habits, but may also relate to current practices. Even in old age, persons of higher educational status report use of more preventive health services than their less educated peers.[20]

Health policy research in aging is frequently directed toward the high cost users of medical care, those who are very ill or those in institutions.[4,7] For health promotion and disease prevention to be most effective in old age, the emphasis may need to shift from the very disabled old to the less disabled population at high risk of developing disability. A recent NIH conference on geriatric assessment highlighted the need for targeted intervention strategies[21]; those with impairments of walking, climbing, stooping, or lifting but without difficulty on ADL or IADL may define such a group with incipient problems. In fact, among those without ADL or IADL problems at baseline in 1984, those non-physically able were at twice the risk of developing a problem with ADL or IADL at follow-up, RO = 2.1 (CI = 1.6, 2.9) compared with the physically able. Our analysis also showed risks of institutionalization, hospital use or death to be greater in this group in comparison to the physically able.

Our criteria were restricted to physical ability because we were interested in defining higher level function to be comparable to Activities of Daily Living and Instrumental Activities of Daily Living, both of which emphasize motor function. There is a need for work to examine and redefine higher level function in old age to include dimensions other than pure physical function.[22]

Despite our intention to choose measures equally applicable to both sexes, women consistently reported more difficulty with each physical activity, the largest difference was observed for difficulty lifting 10 pounds. The interpretation of this difference by sex is unclear but is consistent with other data for the old-old.[23] Whether differences in reported difficulty relate to true differences in physical ability or gender-specific styles of interpretation of these types of questions is unclear.

A potential methodologic problem arises in that a large proportion of persons reported as physically able at baseline no longer met the criteria for physical ability at follow-up and a much smaller percent of those not physically able at baseline were physically able at follow-up. Whether such dramatic changes occur in functional status for populations of very old persons over short periods of time will be clearer as additional follow-up is completed. This finding highlights problems with functional status measures[24,25] as these measures capture not only the "true" measurement of function but variation due to factors such as acute health status and response set. Others have noted that change in functional patterns may be particularly difficult to delineate due to the

subjective nature of the measure.[26] There is a critical need for methodological work to assess measurement of true change in status from other sources of intra-individual variation.

These data suggest a substantial proportion of persons aged 80 or older have no functional impairments and maintain that status over short-term follow-up. Risk factors associated with continued physical ability included weight, arthritis, and cardiovascular disease, factors which may be amenable to intervention. Additional longitudinal follow-up and analysis of medical utilization data will provide the opportunity to continue to examine the relationship of physical ability to morbidity and mortality over time.

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ERRATA

In: the Letters Section of the Journal, February 1989 issue, page 233, an author's affiliation was incorrectly listed. Bradley J. Bradford, MD, is Chairman of the Department of Pediatrics at Mercy Hospital of Pittsburgh, Pennsylvania, and Clinical Assistant Professor of Pediatrics at the University of Pittsburgh School of Medicine, *not* Chairman of the Department of Pediatrics at the University.

* * *

In: Meade CD, Byrd JC: Patient literacy and the readability of smoking education literature. Am J Public Health 1989; 79:204–205. Of the 41 references cited in the text, the last five—numbers 37–41—were inadvertently dropped. They are printed below. The Journal staff regrets the error.

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Predictors of Healthy Aging: Prospective Evidence from the Alameda County Study

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Abstract: Long-term predictors of high levels of physical functioning were examined in a representative sample of Alameda County, California residents followed from 1965 through 1984. The cohort investigated in this study was born between 1895 and 1919, with survivors being age 65 to 89 at the time of follow-up. A scale of physical functioning was developed from a comprehensive set of questionnaire items which assessed the full spectrum of physical functioning. Those scoring in the top 20 percent, defined as healthy aging, were compared to the remainder of the cohort, including those who died and those with lower levels of functioning at follow-up. After adjustment for age and functional status at baseline, the following variables were predictive of high functioning at follow-up 19 years later: race (those not Black), higher family income level, absence of hypertension, absence of arthritis, absence of back pain, being a non-smoker, having normal weight, and consuming moderate amounts of alcohol. Sex did not predict high function because of the counterbalancing effects of higher survival in females but greater likelihood of high functioning among surviving males. (*Am J Public Health* 1989; 79:703–708.)

Introduction

There have been dramatic improvements in life expectancy in this century, and mortality declines in the population age 65 and above have contributed substantially to these gains in the past 40 years. An individual reaching age 65 can now expect to live an average of 17 additional years[1] and life expectancy for older persons is expected to continue to rise into the next century.[2] An important and unresolved issue is whether the postponement of death that results from declining mortality rates will be accompanied by a similar postponement in serious morbidity.[3–6] In epidemiologic studies of aging, therefore, there are enormous potential benefits in identifying factors which not only influence longevity but also promote survival in an independent, highly functional state of health.

Rowe and Kahn have recently proposed use of the term successful aging to describe the concept that decline in organ functioning and physiologic processes are not necessarily an inherent part of aging.[7] Extending this concept for the purpose of identifying the healthy older person is not a straightforward process, however. In one of the few studies which has attempted to operationalize the concept of healthy aging in a population of older persons, Benfante, Reed and Brody considered healthy aging to be present in a subset of the Honolulu Heart Study cohort which was disease-free at baseline and had no incident chronic diseases over a 12-year follow-up period.[8]

The assessment of functional capacity can also serve as a valuable indicator of positive outcomes in aging as well as an indicator of disability. This study examines one aspect of healthy aging, namely the maintenance of very high levels of physical functioning. The study begins in middle age and examines functional status almost two decades later.

Methods

Study Population

The population used in these analyses is a subsample of a study of adults in Alameda County, California, the study design and sampling for which have been reported in detail elsewhere.[9,10] In 1965, a representative sample of adults age 20 and older (16 if ever married) was selected. In total, 6,928 individuals completed an extensive questionnaire about their health and behavioral, social and psychological aspects of their lives. Respondents were again contacted in 1974 and in 1982/83 a random 50 percent sample of 1974 respondents was contacted. In 1984, interviews were conducted with those individuals who were born prior to 1920 and who had completed the 1982/83 questionnaire.

The current analyses are limited to the cohort born between 1895 and 1919, who were age 65 to 89 at follow-up in 1984 or who died prior to that time. To be eligible for the 1984 survey, participants had to have been selected in the 50 percent random sampling of the cohort done for the 1982/83 survey. The analytic cohort for this study includes all those participating in the 1984 survey, all those dying between 1982 and 1984, and a random 50 percent sample of those dying between 1965 and 1982. This cohort consists of 841 individuals, of whom 345 (41.0 percent) died prior to 1984. Of the 496 survivors for whom an interview was completed in 1984, a proxy respondent was required for 31 (6.3 percent) due to a physical or cognitive disability.

In the 1965 baseline sample, 2,392 persons were born between 1895 and 1919. By 1974, 319 (13.3 percent) were known to have died. Of those remaining, 1,677 (80.9 percent) completed the 1974 questionnaire, 346 (16.7 percent) were located but did not respond, and 50 (2.4 percent) were not located. In 1982, 837 of the 1,677 persons responding in 1974 were randomly selected for follow-up. Between 1974 and 1982, 144 deaths occurred in this group (17.2 percent). Of the remaining 693 persons, 565 (81.5 percent) completed the questionnaire, 85 (12.3 percent) were located but did not respond, and 43 (6.2 percent) were not located. By 1984, 27 deaths occurred in the 1982 respondents (4.8 percent). Of the remaining 538 persons, 496 (92.2 percent) completed the 1984 interview, 30 (5.6 percent) were located but not successfully interviewed, and 12 (2.2 percent) were not located.

Those not located and those who were located but refused to participate were different from the cohort analyzed in this study in respect to several baseline variables. Among those born between 1895 and 1919, the group lost to follow-up tended to be younger at baseline than those successfully followed (60.2 percent of the group not located were ages 46–55, compared to 62.2 percent of the located nonrespondents and 51.6 percent of the analytic cohort). Persons lost to follow-up were more likely to be Black (Blacks comprised 19.4 percent of the group not located, 18.8 percent of the located

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nonrespondents, and 14.3 percent of the analytic cohort) and had lower family incomes (those classified as having inadequate or marginal incomes at baseline comprised 36.4 percent of the group not located, 25.8 percent of the located nonrespondents, and 22.5 percent of the analytic cohort).

Outcome Status

Mortality—Deaths were ascertained through a computer-matching procedure with the California Death Registry.[11,12] Additionally, extensive in-state and out-of-state tracing was done in 1974 and 1982 for participants not responding to questionnaires.[9] For previously unascertained deaths which were reported by next-of-kin or others at the time of the 1984 survey, death certificates were requested for confirmation.

Physical Functioning—The 1984 interview contained an extensive section on physical functioning and disability, including basic activities of daily living, instrumental activities of daily living (doing housework and shopping), and more vigorous physical functioning (difficult household tasks, exercise and recreational activities). To assess the full spectrum of functioning and identify those in the cohort functioning at the highest levels, a comprehensive, summative scale was constructed. Selection of appropriate questionnaire items for this scale was accomplished by entering variables representing all items related to physical functioning into a principal factor analysis using oblique rotation.[13,14] A large number of the items loaded on the first two factors (see Appendix), which were the only factors with eigenvalues greater than 1.

Scales of basic functioning and exercise were created using items loading on the first and second factors, respectively. All items in the basic function and exercise scales were then combined to create a summary scale of functioning. While these factors represent different constructs, the scales were combined because of the underlying correlation of the two factors (reference axis correlation = .33) and because of the clear hierarchical relationship of these scales. None of those with poor scores on the basic function scale had high scores on the exercise scale while those with high scores on the exercise scale all had very high scores on the basic function scale. Cronbach's alpha was .90 for the combined scale. The distribution of scores on the combined scale is shown in Figure 1. Scores on this scale ranged from 0 to 92, with a median of 72, a mean of 68.5 and a standard deviation of 15.0. Those scores of 80 and above represent the top quintile of scores. Individuals in this group are classified as having high functioning for the purpose of the analyses to be presented here.

Baseline Variables

Data on demographic factors, chronic conditions, and health-related practices examined in these analyses were obtained from responses to the baseline 1965 questionnaire. Health-related practices examined in these analyses were those previously demonstrated in the Alameda County Study to be associated with increased probability of survival in younger[9] and older[15] persons.

Because not all participants began the study in 1965 with the same level of physical functioning, it was necessary in multivariate analyses to adjust for baseline functional status. Ideally, the same scale used in 1984 would be used to assess 1965 functional status, with this scale score being entered as an independent variable in all analyses. However, the variables making up the 1984 combined function scale were not all collected at baseline. For those variables which were collected in both 1965 and 1984, an abbreviated scale of functioning was empirically developed which, using 1984 data, had a .80 correlation with the full, 1984 combined

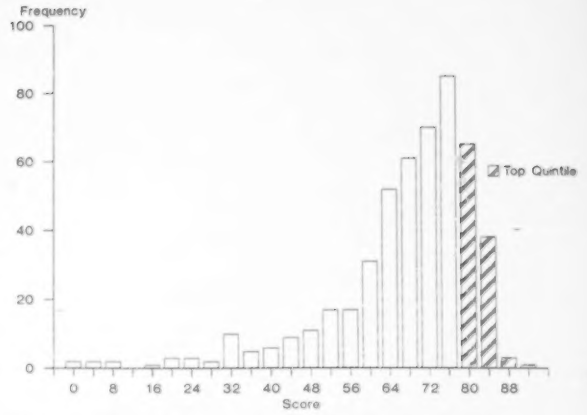

FIGURE 1—Distribution of Scores on the Combined Function Scale

function scale. In practice, this scale, using 1965 data, was entered into models as two variables, one containing basic functioning items and the other containing exercise-related items. In all models which were constructed, these 1965 variables remained highly significant predictors of high functioning in 1984. The inclusion of chronic conditions in the full models to be presented also provides an additional adjustment for health status at baseline.

Statistical Analyses

Analyses compare those in the highest quintile of scores on the combined scale of physical functioning with the remainder of their cohort, including those who died prior to follow-up. Multiple logistic analysis is used to assess the multivariate associations of baseline variables with outcome while simultaneously adjusting for age, race, sex, and baseline functional status. Parameters of the logistic function were determined using the iterative maximum-likelihood technique as described by Walker and Duncan.[16]

In the first set of analyses, the high function group is compared to all other members of the cohort, nearly half of whom had died prior to the follow-up. To investigate the possibility that significant associations found in the first set of analyses resulted simply because of the strong, well-recognized associations of many of the baseline variables with mortality, multiple logistic models using a trichotomous outcome were constructed. This allowed, in one model, for the separate analyses of associations between baseline factors and: 1) the outcome high function versus deceased, and 2) the outcome high function versus moderate to low function. These analyses were accomplished using the CATMOD procedure of SAS, modified for logistic regression.[17] This procedure uses the maximum-likelihood estimation method and yields the same results as the SAS LOGIST procedure when using a dichotomous outcome variable, but has the advantage of being applicable to polychotomous outcomes as well.

Results

Among the 841 individuals in the cohort under study, there were 496 survivors (59.0 percent) at follow-up in 1984. Of these 496 survivors, there were 107 with scores at follow-up of 80 and above on the combined function scale. This high function group comprised 12.7 percent of the original cohort of 841 persons, with low to moderate function survivors comprising 46.3 percent and decedents 41.0 percent.

Results that follow will focus on the relation of baseline states to the outcome of interest, high functioning. For those in the study cohort (including decedents) reporting no high blood pressure at baseline, 14.9 percent went on to be in the high function group 19 years later as compared to only 2.1 percent in the group reporting high blood pressure (Table 1). Those reporting no heart disease, chest pain, shortness of breath, diabetes, arthritis, joint swelling, pain or stiffness and back pain were also significantly more likely to have high functioning at follow-up than those reporting these conditions. There was a low prevalence of self-reported diabetes, stroke, and cancer at baseline, but it is of interest that no one reporting these conditions had high function at follow-up.

For demographic variables, the proportion with high function at follow-up was significantly higher for younger individuals, those not Black, those with very adequate or adequate income, and those married at baseline (Table 2). Among the health practices, non-smoking, having moderate weight, moderate alcohol consumption, and eating breakfast regularly were all significantly associated with high function at follow-up. Physical activity was not studied as an independent variable because many of the items in this scale are incorporated in the scale used to adjust for baseline functional status in multivariate models.

Models are presented in Table 3 which evaluate demographics, chronic conditions, and health practices in three separate analyses, all adjusted for age, race, sex, and baseline functional status. As in the previous results, sex is not predictive of high function. Marital status is unimportant when adjusted for other demographic variables and baseline function, while the importance of age, race, and family income remains. Of the chronic diseases, the reported absence of high blood pressure and arthritis are clear predictors of high function, with back and chest pain being somewhat less so. Being a past or never smoker and having moderate body weight are associated with high function. Those with moderate alcohol intake were 2.4 times more likely to have high function at follow-up than abstainers and 1.7 times more likely to have high function than heavy drinkers. Eating breakfast, hours of sleep, and snacking were not associated with high function in this model. The strengths of association and 95 percent confidence intervals (CI) of all variables entered into a combined model showed little difference from the individual models.

TABLE 1—Percent of Cohort (including decedents) with High Function at Follow-up According to Baseline Chronic Conditions: Alameda County Study

Chronic conditions (past 12 months)		
High blood pressure	No	14.9
	Yes	2.1
Heart disease	No	13.5
	Yes	1.9
Chest pain	No	13.9
	Yes	4.9
Shortness of breath	No	13.8
	Yes	5.9
Swollen ankles	No	13.5
	Yes	9.2
Leg cramps	No	13.4
	Yes	10.3
Stroke	No	12.9
	Yes	0.0
Bronchitis	No	12.6
	Yes	15.1
Asthma	No	12.7
	Yes	16.0
Diabetes	No	13.2
	Yes	0.0
Cancer	No	13.1
	Yes	0.0
Arthritis	No	15.6
	Yes	4.6
Joint swelling/pain/stiffness	No	14.4
	Yes	8.5
Back pain	No	14.5
	Yes	7.8

TABLE 2—Percent of Cohort (including decedents) with High Function at Follow-up According to Baseline Demographic Characteristics and Health Practices: Alameda County Study

Demographic Characteristics		
Age (Baseline/Follow-up)	46–49 / 65–68	20.9
	50–59 / 69–68	15.0
	60–70 / 79–89	4.1
Sex	Male	13.9
	Female	11.7
Race	Not Black	14.3
	Black	3.3
Family Income	Very adequate-adequate*	16.2
	Marginal-inadequate	2.8
Current Marital Status	Married	14.6
	Not married	6.3
Health Practices		
Smoking	Past-Never	15.6
	Current	8.6
Relative Weight	Moderate**	14.4
	Other	6.3
Alcohol Consumption (drinks month)	None	7.0
	1–60	15.1
	>60	11.6
Eating breakfast	Regularly	14.1
	Sometimes-rarely	7.7
Hours of Sleep Per Night	7–8	13.5
	<7 or >8	10.2
Snacking	Rarely-never	14.2
	Sometimes-always	12.1

*Total household income adjusted for household size and scored in comparison to federal standards for 1965.

**Weight within 10% under and 30% over ideal weight according to Metropolitan Life Insurance 1959 standards.

Table 4, based on the trichotomous logistic model, demonstrates the association of baseline variables with three levels of outcome in 1984: high function, low to moderate function, and deceased. For example, those with no high blood pressure at baseline are 6.7 times more likely to end up in the high function group as in the deceased group and 4.1 times more likely to be in the high function group at follow-up compared to the low to moderate function group. Variables shown to be predictors of high function in previous models continue to be predictors of high function when compared both with dying during follow-up and with low to moderate function at follow-up.

Discussion

We used physical functioning as a measure of outcome, a departure from the usual disease-specific outcome in epidemiologic studies. There are numerous theoretical and practical advantages in characterizing older persons according to their level of functioning.[18] While it is certainly desirable to age with no chronic diseases, it can be problematic to characterize healthy aging according to diagnosed diseases. How important or severe must a disease be to remove someone from the healthy aging group? How intensively must we search for disease in order to declare someone disease-free? It is common for individuals who have chronic diseases but who are well compensated to live long, active, and vigorous lives, an outcome to which we all aspire. It is this outcome which is addressed in these analyses.

TABLE 3—Multiple Logistic Models Showing Association between Demographic Characteristics, Chronic Conditions and Health Practices and High Function at 19 Years: Alameda County Study

		Odds Ratio (95% CI) Demographic Model*	*Odds Ratio (95% CI)** Combined Model*
Age (10 years)	Younger/Older	2.3 (1.6, 3.5)	2.9 (1.9, 4.6)
Sex	Male/Female	1.2 (0.7, 1.9)	1.2 (0.7, 2.0)
Race	Not-black/Black	3.4 (1.2, 9.8)	2.9 (1.0, 8.7)
Family Income	Very adequate- adequate/ Marginal-inadequate	4.0 (1.5, 10.8)	4.0 (1.4, 11.2)
Current Marital Status	Married/Not married	1.3 (0.6, 2.7)	
		Chronic Condition Model	
High Blood Pressure	No/Yes	4.9 (1.4, 16.7)	4.3 (1.2, 15.1)
Heart Disease	No/Yes	1.1 (0.1, 10.2)	
Chest Pain	No/Yes	2.3 (0.7, 7.5)	
Shortness of Breath	No/Yes	1.1 (0.5, 2.8)	
Arthritis	No/Yes	2.6 (1.1, 6.0)	2.8 (1.3, 6.0)
Joint Swelling/Pain/Stiffness	No/Yes	1.0 (0.5, 1.9)	
Back Pain	No/Yes	1.9 (1.0, 3.6)	2.0 (1.1, 3.8)
		Health Practices Model	
Smoking	Past-Never/Current	2.8 (1.6, 4.8)	3.0 (1.8, 5.1)
Relative Weight	Moderate/Other	2.4 (1.1, 5.2)	2.3 (1.1, 4.9)
Alcohol/Consumption	1–60/None	2.4 (1.2, 4.8)	2.1 (1.1, 4.1)
(drinks/month)	1–60/>60	1.7 (0.6, 5.0)	1.7 (0.6, 5.1)
Eating Breakfast	Regularly/Sometimes-rarely	1.5 (0.7, 3.1)	
Hours of Sleep	7–8/<7 or >8	0.9 (0.5, 1.7)	
Snacking	Rarely-never/ Sometimes-always	1.3 (0.8, 2.1)	

*Demographic, chronic condition and health practices models: odds ratios adjusted for age, race, sex, baseline functional status and all variables in the model.

**Combined model: odds ratios adjusted for baseline functional status and all variables in the model.

In choosing an outcome of functioning rather than disease status, it is important to recognize that predictors of functional status identified in this study have their influence on function through specific diseases as well as patterns of accommodation and adaptation to these diseases. However, there are important gaps in our understanding of these pathways from disease to disability in older persons.[19] Using functioning as an outcome measure does not deny the importance of these pathways, it simply uses function as an overall measure of the impact in the older person of multiple diseases, conditions, and environmental influences.

The outcome measure in this study was a summative scale created from a large group of questionnaire items representing activities with a wide spectrum of difficulty. It was created to identify participants who were not only non-disabled but were also functioning at a high level. Nevertheless, it has a number of limitations. It assesses only physical functioning and does not take into account the domains of cognitive or psychological functioning directly. Many of the items querying activities of daily living and more difficult tasks have a large cognitive component, however, and it is likely that an individual with moderate to severe dementia would have difficulty with at least some of these activities and not score in the top quintile. The distribution of scores on this scale was unimodal, so the classification of the top quintile as representing high function was somewhat arbitrary. Scores on a summative scale, as opposed to a hierarchical scale, can be obtained in a number of different ways and it can be difficult to interpret what a particular score means. However, the construction of the scale as it was done here allows for a good characterization of those classified as having high function. To obtain a score in the top quintile, individuals had perfect or near perfect scores on the basic function component of the combined scale and reported engaging in at least some of the more vigorous activities on the exercise component.

The difficulty in tapping the high end of the functional spectrum using questionnaire items may have led to misclassification of individuals on the outcome measure used here. Those with no opportunity or inclination to engage in more vigorous activities, but who might have the capacity to perform them, would have scores below the top quintile and would not be classified as having high functioning. This potential misclassification could be a source of bias affecting our findings that both income and race are strongly associated with high function. To evaluate this, an analysis was performed using the basic function scale alone as the outcome. Employing this scale as a continuous variable and studying survivors, both income and race were found to be strongly associated with this outcome, even after adjusting for all other significant variables in a multiple linear regression model.

A number of other factors known to be associated with all-cause mortality and specific diseases were demonstrated in this study to be predictive of high function. These include smoking status, relative weight, alcohol consumption, and blood pressure. Physical activity has traditionally been one of the seven health behaviors studied in the Alameda County cohort, but it cannot be studied here as a predictive variable because it is part of both the outcome variable and the baseline adjustment variable. In all multivariate models, the baseline exercise variable is significantly associated with high function at outcome. While it is not correct to infer that it is causally related to high functioning, its association with high functioning offers evidence that those with high levels of physical activity at baseline continue to have high physical activity at follow-up. Arthritis and back pain—factors usually associated with disability but not mortality—remain significant predictors of high function after adjustment for many other chronic conditions, demographics, and health practices. Interestingly, in the logistic model with a trichotomous outcome, the absence of arthritis and back pain predict survival with high function versus mortality as well as

TABLE 4—Multiple Logistic Models Showing Association between Demographic Characteristics, Chronic Conditions and Health Practices and 19-Year Status (High function, Low/moderate function, deceased)*

		High function vs Deceased	High function vs Low/moderate function
		Odds Ratio (95% CI)	Odds Ratio (95% CI)
Demographic Model			
Age (10 years)	Younger/Older	5.1 (3.2, 8.0)	1.5 (1.0, 2.3)
Sex	Male/Female	0.5 (0.3, 0.8)	1.2** (0.7, 1.9)
Race	Not-black/Black	4.1 (1.4, 12.4)	3.1 (1.1, 9.3)
Family Income	Very adequate-adequate/ Marginal-inadequate	4.9 (3.0, 8.2)	3.1 (1.4, 10.2)
Chronic Conditions Model			
High Blood Pressure	No/Yes	6.7 (1.9, 23.8)	4.1** (1.2, 14.1)
Arthritis	No/Yes	2.4 (1.1, 5.3)	3.0 (1.4, 6.5)
Back Pain	No/Yes	1.9 (1.0, 3.7)	2.0 (1.1, 3.6)
Health Practices Model			
Smoking	Past-Never/Current	6.1 (3.3, 8.3)	2.2** (1.3, 3.8)
Relative Weight	Moderate/Other	2.6 (1.1, 6.1)	2.4 (1.1, 5.1)
Alcohol Consumption	1–60/None	3.1 (1.5, 6.5)	2.2 (1.1, 4.4)
	1–60/>60	2.5 (0.7, 8.2)	1.3 (0.4, 4.0)

*Odds ratios adjusted for age, race, sex, baseline functional status and all variables in model.
**For this variable, the odds ratio for high function vs low/moderate function is significantly different from the odds ratio for high function vs deceased at $p < .05$.

survival with high function versus low to moderate function.

An important finding in this study is that in analyses in which the high function outcome group is compared to the remainder of the cohort, sex is not a predictor of high function; the logistic analysis with trichotomous outcome suggests the reasons: females are more likely to be in the high function group when comparing this group with decedents, but among survivors males have a greater likelihood of being highly functional. Sex did not predict function at older ages because of the poorer survival in males and the poorer functional outcome in surviving females compared to surviving males. This is compatible with the findings of Katz and colleagues who found, in a cohort of persons already age 65 and older, that differences in average remaining disability-free years between males and females in different age groups were quite small and did not reflect the large differences in remaining life expectancy at these ages.[20]

There are no other prospective studies with long-term follow-up which provide data on how factors assessed in a cohort which is middle-aged at baseline predict high function when the survivors have passed age 65. In their study of factors associated with remaining disease-free in Japanese men ages 46–69 at baseline, Benfante, Reed and Brody found that systolic blood pressure, total cigarettes smoked, body mass index, and alcohol consumption measured at baseline were all inversely related to remaining disease-free at follow-up 12 years later.[8] In analyzing the effect of alcohol intake, they found the highest probability of remaining disease-free among alcohol abstainers, in contrast to our finding that moderate drinkers had a higher probability than abstainers of having high function at follow-up. Their physical activity index was not predictive of remaining disease-free.

Several studies have explored the relation of risk factors to the development of disability. Pinsky and colleagues used the Framingham Heart Study to investigate how traditional risk factors for heart disease predict disability in a subsample of survivors who remained free of diagnosed cardiovascular disease.[21] Ever having hypertension was predictive of disability in both men and women and elevated body mass index was predictive in women, but cigarette smoking was not associated with disability. Palmore studied demographic predictors of disability in a cohort followed for approximately 10 years, whose surviving members were age 72 and older at follow-up.[22] Adjusting for baseline disability, only age and a scale of physical health (measured in terms of disease presence and severity) were significant predictors of disability at follow-up. Sex, race, marital status, education, and income were not independent predictors of disability. Branch investigated incident disability in a cohort of persons age 66 and older who were followed for five years.[23] A number of health practice variables similar to those used in the Alameda County Study were evaluated. Slowed down activity in women and current or past smoking in men were the only variables predictive of disability in this study. It is of interest that the latter two studies, in which cohorts were already at older ages at baseline, found few variables predictive of disability.

One limitation of the present study is the selective loss during the follow-up period of more persons in the younger part of the study population, more Blacks, and more individuals with low family income levels. These losses are likely to bias the estimation of overall levels of physical functioning in 1984. However, specific risk estimates would be biased only if the associations of baseline risk factors with the outcome were different for those lost to follow-up and those remaining in the cohort. Because individuals who are lost to follow-up tend to have poorer outcomes than those who remain, it is likely that the inclusion in the study of those Blacks and low income individuals who were lost would have resulted in even stronger associations of race and income with functional outcome.

Findings in this study should also be interpreted in light of the potential changes in the baseline variables during the long follow-up period. In certain cases it is advantageous to measure risk factors during middle age. Assessing income at older ages may not give a valid picture of income during the pre-retirement years, so our baseline measure of income has particular advantages. However, some variables likely to change significantly after their baseline assessment may not reflect status during majority of the follow-up period. As an example, marital status as reported at baseline, which was not a significant predictor of high functioning, may not have been a good measure of marital

status during the 19 years of follow-up. In an older population such as this, many individuals married at baseline may have been widowed in the first few years of follow-up.

It is also doubtful that a single report of high blood pressure at baseline was representative of blood pressure status during the long follow-up period. Misrepresentation of blood pressure status could result from unreliable reporting, the use of antihypertensive medications, which was not assessed at baseline, the degree of blood pressure control and the development of hypertension after the baseline assessment. This type of misclassification would lead to a conservative bias in assessing the association of blood pressure with high function. Indeed, it is striking that a single question about high blood pressure asked 19 years prior to assessing function was so highly predictive of high function. These findings, along with those of Benfante and colleagues and Pinsky and colleagues, indicate that absence of hypertension may be one of the most important predictors of healthy aging. The benefits of hypertension control on cardiovascular disease are well documented.[24–26] Studies to assess how hypertension control affects overall health status and disability in older persons have not been done and would be very valuable.

As mortality rates decline and the average age at death is postponed, there will be a marked expansion in overall population morbidity levels if the onset of disease and disability is not postponed an equal or greater number of years. It is clear that in the next century the older population will survive longer and grow to unprecedented size. The findings from this study indicate the possibility that active intervention in areas which promote healthy aging might lead to an aging population which retains high levels of function for a longer proportion of their lives and is therefore less dependent on their families and the health care system.

APPENDIX

This list summarizes items loading on factors 1 and 2. The first factor contains items which mostly relate to basic functioning, that is, activities necessary in day-to-day life. Loading heavily on the second factor were items related to exercise. Miscellaneous items loaded on a third factor and were dropped from further scale construction. The grouping pattern of items on the first two factors was understandable and in general reflected what one might have constructed *a priori*. These two scales were therefore retained and items were given equal weighting in the creation of summary measures.

Factor 1—Basic Function Scale

Activities of daily living: bathing, grooming, dressing, eating, transferring from bed to chair, using the toilet
Walking across a small room
Climbing stairs
Walking one-half mile
Pushing or pulling a large object
Stooping, crouching, kneeling
Lifting and carrying weights over 10 pounds
Gardening
Meal preparation
Heavy housework
Difficulty getting places
Shopping

Factor 2—Exercise Scale

Any sports exercise
Walking for exercise
Exercising long enough to sweat
Calisthenics or stretching to exercise
Swimming
Vigorous exercise: jogging, bicycling, aerobics, tennis, dancing, hiking

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Is Early Natural Menopause a Biologic Marker of Health and Aging?

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Abstract: The relation between age at natural menopause and all-cause mortality was investigated in a sample of 5,287 White women, ages 55 to 100 years, naturally-postmenopausal, Seventh-day Adventists who had completed mailed questionnaires in 1976. The age-adjusted odds ratio of death during 1976–82 in women with natural menopause before age 40 was 1.95 (95% confidence interval = 1.24, 3.07), compared to the reference group of women reporting natural menopause at ages 50 to 54. Corresponding odds ratios of death were 1.39 (95% CI = 1.06, 1.81) for natural menopause at ages 40 to 44, and 1.03 (95% CI = 0.84, 1.25) for natural menopause at ages 45 to 49. Among 3,166 White, 55- to 100-year-old, surgically-postmenopausal, Adventist women, there was no relation between age at surgical menopause and mortality. Logistic regression analyses indicated that findings from this study were apparently not due to confounding by smoking, over- or underweight, reproductive history, or replacement estrogen use. (*Am J Public Health* 1989; 79: 709–714.)

Introduction

Advancing age is associated with substantial adverse changes in many physiologic characteristics,[1–10] increases in morbidity,[11] decreases in the capacity for self care at older ages,[12] and increases in mortality.[13] Reproductive functions also show a strong and predictable change with advancing adult age. As age increases between age 30 and menopause, women experience more anovulatory cycles, infertility, chromosomal abnormalities related to oogenesis, and spontaneous abortions.[14] While other organs and physiologic systems usually decline in performance yet continue to function with advancing age, the ovaries decline and essentially cease functioning at menopause.

Several animal species, such as some salmon, mollusk, spiders, and octupus die within a predictable time after mating or reproducing.[15] While these may be extreme examples, they suggest that neural and endocrine mechanisms play a role in determining length of life in some animals. Furthermore, environmental factors that affect the health and aging of reproductive organs also might simultaneously affect other organs and tissues. For example, cigarette smoking adversely affects the function of the reproductive system[16–18] as well as the function of the cardiovascular system, in addition to reducing life expectancy.[18,19]

Based on these considerations, we hypothesized that age at natural menopause would be a biologic marker of the general health and aging state of the individual. Because mortality is an indicator of health and aging, we tested the hypothesis that age at natural menopause is negatively associated with all-cause mortality.

Methods

Study Population and Questionnaires

In 1974, Seventh-day Adventist households in California were identified from church records and a Census Questionnaire mailed to virtually every household. A total of 25,186 White Adventist women 25 to 100 years of age, not of Hispanic origin, were identified by the census. In 1976, Adventists identified in the previous census were mailed another questionnaire called the Lifestyle Questionnaire. Of the 25,186 women originally identified in the census, 19,586 (78 percent) completed the Lifestyle Questionnaire. Variables from the Lifestyle Questionnaire used in the multivariate analysis and our method of coding are described in the Appendix.

Of the 19,586 women 25 to 100 years of age who completed the Lifestyle Questionnaire, 10,198 were 55 to 100 years of age. Of these women, 5,287 (51.8 percent) reported they were naturally-postmenopausal, 3,166 (31.0 percent) surgically-postmenopausal, 75 (0.7 percent) pre-menopausal, and 1,670 (16.4 percent) were either unknown for type or age at menopause or had an implausible age at menopause (i.e., age at menopause was greater than 64 or greater than the subject's age). While there is no generally accepted normal range for age at menopause,[20] age 64 was used in this study as the last acceptable year. Thus, the study population for this report was the 5,287 naturally- and 3,166 surgically-postmenopausal women who were 55 to 100 years of age at the start of the study (i.e., restricted to those ages where virtually all women were postmenopausal).

Mortality Surveillance

Mortality data in this report cover the time between the completion of the Lifestyle Questionnaire in 1976 and the end of the follow-up on December 31, 1982. Of the 19,586 women who completed the Lifestyle Questionnaire, 17,447 (89.1 percent) were known to be alive when the follow-up ended because they completed a mailed questionnaire in 1983, a death certificate indicated they died after 1982, or a personal contact by phone indicated they were alive in 1983. A total of 1,439 (7.3 percent) of the 19,586 women died during the 1976–82 time period. All deaths were confirmed by death certificates. Vital status at the end of follow-up in 1982 could not be determined for 700 (3.6 percent) of the 19,586 women. Data for these 700 women were compared to data on the computerized California state death certificate files for 1976–82 and the National Death Index for 1979–83 but no positive matches with an accompanying death certificate were located. Thus, these 700 women were treated as "alive" in the data analyses.

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Statistical Analyses

Prior to analyses, age at menopause was categorized into five groups (i.e., less than 40 years, 40 to 44, 45 to 49, 50 to 54, and 55 to 64) with those having menopause at ages 50 to 54 as the reference category (for all odds ratios). Age-specific odds ratios and their confidence intervals were estimated by methods described by Mantel and Haenszel.[21] Age-adjusted odds ratios and their precision-based confidence intervals were estimated by methods for stratified data described by Woolf and Haldane.[22,23] Odds ratios adjusted for other variables were estimated by a maximum likelihood logistic regression model.[24]

Results

Age at Natural Menopause

These analyses were based on 5,287 naturally-postmenopausal women who were 55 to 100 years of age at the start of the study in 1976. The mean age at natural menopause in these women was 49.21 (SE = 9.70); mode = 50, median = 50, 25th percentile = 46, and 75th percentile = 52.

Of these 5,287 women, 772 died during the study (i.e., 1976–82), 4,356 survived to the end of the study, and 159 were presumed to be alive at the end of the study (i.e., we lost contact with these subjects sometime during the study). However, there was no difference in the age at natural menopause between the 4,356 women known to have survived and the 159 presumed to have survived to the end of the study. Age-specific Spearman rank correlations ranged from −0.01 to 0.07 for the association between age at natural menopause and "living status" (i.e., 0 = presumed alive; 1 = known alive).

Age-specific results (Table 1) indicated that the odds ratio of death decreased with increasing age at natural menopause until age 55. This negative association between age at natural menopause and death was evident within each age group except women 85 years of age and older in 1976. Other age-specific results (Figure 1) also indicated similar negative associations between age at natural menopause and the probability of death (denominator based on counts of people) and the rate of death (denominator based on person-years of observation). Because of these findings, all other analyses were limited to odds ratios (based on counts of people).

To control for potential confounders, logistic regression analyses were done using four models which adjusted for (I)

TABLE 1—Age at Natural Menopause and the Age-specific and Age-adjusted Odds Ratios of Death in 5,287 California Seventh-day Adventists, 1976–82

Age in 1976	Age at Natural Menopause (years)	Odds Ratio of Death[a] (95% CI)	Number At-Risk (1976)	Number Deaths (1976–82)	Number Person-Years[b] (1976–82)
55–64	<40	2.03 (0.46, 8.91)	31	2	195
	40–44	1.55 (0.70, 3.45)	160	8	1006
	45–49	1.50 (0.86, 2.59)	497	24	3126
	50–54	1.00	914	30	5749
	55+	1.57 (0.85, 2.92)	316	16	1988
		trend p = 0.51[p = 0.10][c]			
65–74	<40	2.54 (1.21, 5.30)	48	10	302
	40–44	1.55 (0.96, 2.50)	188	26	1183
	45–49	1.00 (0.68, 1.45)	534	50	3359
	50–54	1.00	777	73	4887
	55+	0.81 (0.49, 1.34)	283	22	1780
		trend p = 0.006[p = 0.01]			
75–84	<40	1.67 (0.75, 3.71)	28	10	176
	40–44	1.26 (0.83, 1.91)	139	41	874
	45–49	1.05 (0.77, 1.44)	373	97	2346
	50–54	1.00	476	119	2994
	55+	0.91 (0.58, 1.44)	129	30	811
		trend p = 0.11[p = 0.16]			
85+	<40	1.52 (0.57, 4.07)	19	12	120
	40–44	1.33 (0.69, 2.55)	50	30	315
	45–49	0.76 (0.47, 1.22)	128	59	805
	50–54	1.00	147	78	925
	55+	2.06 (1.04, 4.10)	50	35	315
		trend p = 0.45[p = 0.41]			
Age-Adjusted[d] (Ages 55+)	<40	1.95 (1.24, 3.07)	126	34	793
	40–44	1.39 (1.06, 1.81)	537	105	3378
	45–49	1.03 (0.84, 1.25)	1532	230	9636
	50–54	1.00	2314	300	14555
	55+	1.11 (0.85, 1.46)	778	103	4894
		trend p = 0.02[p = 0.002]			

[a]The reference category for the odds ratio was the group of women that became naturally-postmenopausal at ages 50 to 54.

[b]Women who survived to the end of the study had 6.29 person-years of observation (i.e., the years that elapsed between the beginning of the study on September 15, 1976 and the end of the study on December 31, 1982). For women who died during the study, person-years corresponded to the years that elapsed between the beginning of the study and the date of death.

[c]Trend p refers to the two-tailed p-value for the test of the linear trend in odds ratios across all "age at menopause" groups. The value in brackets refers to the trend p-value across all "age at menopause" groups except the "55+ age at menopause" group.

[d]Age-adjusted by stratification.

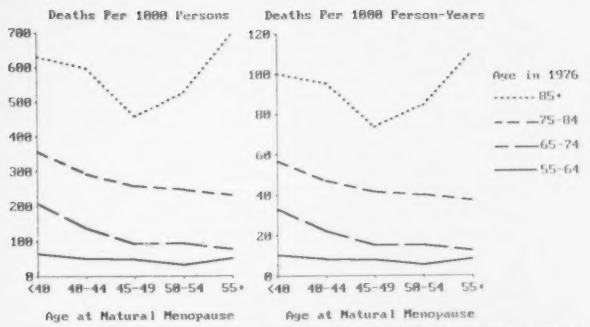

FIGURE 1—Age at Natural Menopause and the Probability and Rate of Death for 5,287 California Seventh-day Adventists, 1976–82

age in 1976, (II) age in 1976 and lifestyle-related factors (III) age in 1976, lifestyle-related factors, and reproductive-related factors, and (IV) age in 1976, lifestyle- and reproductive-related factors, and replacement estrogen use. Variables included in these models and our method of coding are described in the Appendix. Odds ratios of death for different age at natural menopause groups were essentially identical when these four types of analyses were done separately (Table 2). Since the type of model used in the regression analyses did not materially affect these and other odds ratios, all subsequent multivariate-adjusted findings refer to the results from Model IV. Model IV results indicated that women with natural menopause before age 40 had a multivariate-adjusted odds ratio of death of 3.03 (95% CI = 1.62, 5.65), compared to women with natural menopause at ages 50 to 54. The odds ratio of death also decreased in a stepwise fashion with increasing age at natural menopause until age 55.

These odds ratios (Table 2) were larger than those from the age-adjusted stratified analyses (Table 1) because the regression was limited to 3,334 women who had known values for all the variables included in the logistic model rather than the 5,287 women who only had known values for age in 1976 and menopausal status. This increase in the odds ratios appeared to be due, in part, to a selective loss of older subjects: 74 percent (1415/1918) of the 55 to 64 year-old age group were in both analyses, compared to 60 percent (1096/1830) in the 65 to 74 year-old age group, 52 percent (596/1145) in the 75 to 84 year-old age group, and 58 percent (227/394) in the group that was 85 years of age and older. Within each of these age groups, however, women with natural menopause before age 40 had the highest multivariate-adjusted odds ratio of death (ranging from 1.56 among women who were 75 to 84 years of age in 1976 to 8.49 among women who were 85 years of age and older).

In separate analyses to estimate possible effects of replacement estrogens, the odds ratio of death for women who were naturally-postmenopausal before age 40 was 3.33 (95% CI = 1.14, 9.72) within the group that ever used replacement estrogens and 2.91 (95% CI = 1.33, 6.36) within the group that never used replacement estrogens (compared to the reference group who were naturally-postmenopausal at age 50 to 54). Among all naturally-postmenopausal women, the multivariate-adjusted odds ratio of death was 1.13 (95% CI = 0.88, 1.45) for those who ever used replacement estrogens compared to those who never used replacement

TABLE 2—Age at Natural Menopause and the Age-adjusted and Multivariate-adjusted Odds Ratio of Death in 3,334 California Seventh-day Adventists, 1976–82[a]

Model	Factors Adjusted for in Logistic Regression Analyses	Age at Natural Menopause (years)	Multivariate-Adjusted Odds Ratio of Death (95% CI)
I	Age in 1976[b]	<40	3.06 (1.67, 5.62)
		40–44	1.54 (1.06, 2.23)
		45–49	1.10 (0.83, 1.44)
		50–54	1.00
		55+	1.41 (1.00, 1.97)
II	Age in 1976; Lifestyle Factors[c]	<40	2.97 (1.60, 5.50)
		40–44	1.50 (1.04, 2.18)
		45–49	1.07 (0.81, 1.41)
		50–54	1.00
		55+	1.42 (1.01, 2.01)
III	Age in 1976; Lifestyle Factors; Reproductive Factors[c]	<40	3.02 (1.62, 5.62)
		40–44	1.51 (1.04, 2.19)
		45–49	1.07 (0.81, 1.42)
		50–54	1.00
		55+	1.41 (1.00, 1.99)
IV	Age in 1976; Lifestyle Factors; Reproductive Factors; Replacement Estrogens	<40	3.03 (1.62, 5.65)
		40–44	1.51 (1.04, 2.20)
		45–49	1.07 (0.81, 1.42)
		50–54	1.00
		55+	1.40 (0.99, 1.97)

[a]There were 5,287 women with known responses to the menstrual-related questions on the questionnaire. Of these women, 3,334 had known responses to all the lifestyle- and reproductive-related variables included in the multivariate-adjusted analyses (i.e., logistic regression).

[b]For the 3,334 women (known for all the variables included in the multivariate analyses), the age-adjusted (by stratification) odds ratio of death was 3.34 (95% CI = 1.79, 6.21) for natural menopause before age 40, 1.58 (95% CI = 1.08, 2.30) for menopause at ages 40 to 44, 1.12 (95% CI = 0.86, 1.47) for menopause at ages 45 to 49, 1.00 for menopause at ages 50 to 54, and 1.41 (95% CI = 0.99, 2.02) for natural menopause at ages 55 and older.

[c]Lifestyle and reproductive variables are descrioed in the Appendix.

estrogens. Within each age at natural menopause group, ever use of replacement estrogens was associated with only a slight elevation of the odds ratio. Thus, replacement estrogen use did not weaken the association between age at natural menopause and mortality.

Age-adjusted and multivariate-adjusted cause-specific analyses were done in order to determine which of the major causes of death might be responsible for the relationship between age at natural menopause and all-cause mortality (e.g., the multivariate-adjusted analyses on heart disease had the dependent variable equal to "1" if the person died of coronary heart disease and "0" if the person was alive at the end of the study or died of some other cause). Coronary heart disease and stroke mortality each had elevated age- and multivariate-adjusted odds ratios associated with early natural menopause, while the multivariate-adjusted odds of cancer mortality was slightly lower in women with early natural menopause (Tables 3 and 4). Early natural menopause was associated with an excess mortality due to all other causes of death combined. Separate analyses on injuries and other specific causes of death were not performed because there were not enough deaths due to these causes.

Age at Surgical Menopause

Age at surgical menopause was weakly and inconsistently positively associated with all-cause mortality in the age-adjusted analysis on 3,166 women, and the multivariate-adjusted analyses on 2,096 women (who had known responses to all the variables included in the multivariate analyses). The multivariate-adjusted odds ratio of death was 0.75 (95% CI = 0.47, 1.20) in women who were surgically-postmenopausal before age 40, 0.84 (95% CI = 0.52, 1.35) in those surgically-postmenopausal at ages 40 to 44, 0.77 (95% CI = 0.47, 1.28) in those postmenopausal at ages 45 to 49, and 0.89 (95% CI = 0.32, 2.43) in those postmenopausal at age 55 and older, when compared to those surgically-postmenopausal at ages 50 to 54. Age at surgical menopause was unrelated to mortality due to coronary heart disease, stroke, cancer, and all other causes combined.

Replacement estrogen use was associated with a slight decrease in the risk of all-cause mortality in the surgically-postmenopausal women. Among all surgically-postmenopausal women, the multivariate-adjusted odds ratio of death was 0.80 (95% CI = 0.56, 1.14) for those who ever used replacement estrogens compared to those who never used replacement estrogens. These odds ratios for ever use of replacement estrogens were lowest among women with early surgical menopause. For example, within the group that were surgically-postmenopausal before age 40, the multivariate-adjusted odds ratio of death for ever use of replacement estrogens was 0.57 (95% CI = 0.29, 1.10), compared to never use of replacement estrogens.

Discussion

Findings from this study suggest that early age at natural menopause is associated with a substantial excess mortality. In addition, women who report natural menopause at ages 55

TABLE 3—Age at Natural Menopause and the Age-adjusted Odds Ratio of Death due to Selected Causes in 5,287 California Seventh-day Adventists, 1976–82[a]

Cause of Death[b]	Age at Natural Menopause (years)	Age-Adjusted Odds Ratio of Death[c] (95% CI)	Number of Deaths
Coronary Heart Disease	<40	1.29 (0.61, 2.74)	8
	40–44	1.43 (0.95, 2.16)	36
	45–49	1.06 (0.76, 1.46)	74
	50–54	1.00	91
	55+	1.17 (0.76, 1.81)	33
		trend p = 0.42[p = 0.18][d]	
Stroke	<40	3.07 (1.34, 7.03)	8
	40–44	1.39 (0.76, 2.53)	16
	45–49	0.92 (0.56, 1.51)	30
	50–54	1.00	41
	55+	1.32 (0.72, 2.44)	16
		trend p = 0.27[p = 0.06]	
Cancer	<40	1.83 (0.73, 4.59)	4
	40–44	1.20 (0.71, 2.05)	19
	45–49	1.09 (0.74, 1.60)	50
	50–54	1.00	66
	55+	1.07 (0.66, 1.74)	23
		trend p = 0.67[p = 0.60]	
All Other Causes of Death	<40	2.14 (1.15, 3.99)	14
	40–44	1.22 (0.81, 1.84)	34
	45–49	0.99 (0.72, 1.36)	76
	50–54	1.00	102
	55+	0.97 (0.63, 1.50)	31
		trend p = 0.08[p = 0.07]	

[a]See Table 2 for explanation.
[b]Cause of death is the underlying cause of death from the death certificate that was coded by the California State nosologist. Cause-of-death codes from the International Classification of Disease, version nine, were used to classify deaths by major cause; coronary heart disease = codes 410–414, stroke = 430–438, cancer = 140–239, and all other causes of death combined = all other codes.
[c]All results in this table have been age-adjusted by stratification. The reference category for the odds ratio was the group of women that became naturally-postmenopausal at ages 50 to 54.
[d]See footnote "c" in Table 1.

TABLE 4—Age at Natural Menopause and the Multivariate-adjusted Odds Ratio of Death due to Selected Causes in 3,334 California Seventh-day Adventists, 1976–82[a]

Cause of Death[b]	Age at Natural Menopause (years)	Multivariate-Adjusted[c] Odds Ratio of Death (95% CI)
Coronary Heart Disease	<40	1.59 (0.58, 4.40)
	40–44	1.68 (0.96, 2.94)
	45–49	1.02 (0.65, 1.60)
	50–54	1.00
	55+	1.17 (0.67, 2.06)
Stroke	<40	1.87 (0.51, 6.92)
	40–44	1.41 (0.65, 3.07)
	45–49	0.79 (0.42, 1.50)
	50–54	1.00
	55+	1.09 (0.47, 2.50)
Cancer	<40	0.93 (0.22, 4.01)
	40–44	0.83 (0.37, 1.89)
	45–49	1.18 (0.72, 1.94)
	50–54	1.00
	55+	1.35 (0.75, 2.45)
All Other Causes	<40	4.58 (2.10, 9.99)
	40–44	1.56 (0.87, 2.80)
	45–49	1.14 (0.73, 1.79)
	50–54	1.00
	55+	1.43 (0.83, 2.48)

[a]See Table 2 for explanation.
[b]See Table 3 for explanation.
[c]These logistic regression analyses adjusted for age in 1976, replacement estrogen use, and the lifestyle- and reproductive-related variables described in the Appendix.

and older may have a slightly higher risk of death than women with natural menopause at ages 50 to 54. While we have no data to test the hypothesis directly, some of the women who reported menstrual periods at such a late age may have mistaken disease-related vaginal or uterine bleeding for normal menstrual bleeding. If this occurred, it could explain the slightly higher risk of death in these women.

Our study used self-reports of type of menopause and age at menopause from a mailed questionnaire. No studies have investigated the validity of self-reported age at natural menopause. However, a study of 6,591, 30- to 55-year-old nurses found that for questionnaires mailed two years apart, 95 percent of the surgically-postmenopausal women and 82 percent of the naturally-postmenopausal women reported their age at menopause within one year on both questionnaires.[25]

The possibility that our findings were due to confounding by potential correlates of age at menopause was investigated by logistic regression analyses. Controlling for self-reported cigarette smoking, over- and underweight, reproductive history, and replacement estrogen use did not materially alter our findings. Moreover, cigarette smoking is not an important issue in this study population because at the beginning of our study only 1.5 percent of the women were current smokers and 13.2 percent former smokers.

While our findings seem solid, it would be reassuring to find evidence of these associations in other populations. Other studies have investigated only the relationships between age at menopause and the risk of specific diseases. Age at natural or surgical menopause was positively associated with the risk of breast cancer in several studies.[26] Age at menopause (both surgical and natural combined), however, has been negatively associated with osteoporosis.[27,28] Results from two small Swedish studies also suggest that age at menopause is negatively associated with the risk of developing myocardial infarction.[29,30] Among 45- to 54-year-old women in the Framingham Study, naturally- and surgically-postmenopausal women both had 2.7 times the incidence of coronary heart disease (including angina pectoris) compared to pre-menopausal women of the same age.[31]

We found no evidence that estrogen depletion was related to mortality. Replacement estrogen use was not associated with a reduction in mortality in women with early or late age at natural menopause. In spite of the "estrogen depletion" in the group of surgically-postmenopausal women in our study, age at menopause was unrelated to mortality in these women.

Premature natural menopause may be related to mortality because of aging-related physiologic processes that are associated with early natural menopause. The function of the ovaries throughout life and the start and permanent cessation of menstruation are highly dependent on age. The average number of ovarian germ cells or oocytes present in human females appears to be greatest around the fifth or sixth month of fetal development.[32] Thereafter, the number of surviving oocytes regularly declines pre- and post-natally until menopause is apparently precipitated by an insufficient number of surviving oocytes and follicles.[14] This involution of the ovaries would appear to be a natural consequence of aging (as opposed to disease), particularly since the average age at menopause appears to have been remarkably consistent in western countries throughout the last three decades.[33]

Ovarian aging, as reflected by age at natural menopause, may be highly correlated to the aging of other tissues and to the mortality produced by aging (i.e., the increase in mortality with age that would probably inevitably occur even in the absence of disease). Further investigation is needed to pinpoint whether age at natural menopause is only related to a few diseases, or whether it is a marker of multiple aspects of health and aging.

APPENDIX

Variables from the Lifestyle Questionnaire

The Lifestyle Questionnaire contained questions on food and beverage consumption, physical activity, smoking, occupation, psychosocial characteristics, and medical history. The questionnaire also contained questions on whether the subject's menstrual periods had permanently stopped, why they stopped, and the age of the subject when they stopped. Each participant's responses to six questions on the Lifestyle Questionnaire also were used to create a summary index of participation in church activities (e.g., attending church services, holding church office, and attending church social events). Variables used in the regression models (and our method of coding) are described below:

Age In 1976

1. chronologic age (as one continuous term).

Lifestyle Factors

1. cigarette smoking status in 1976 (as a dummy, binary, term indicating whether they currently smoked),
2. pack-years of cigarettes smoked in the past (as one continuous term),
3. Quetelets Index (using dummy terms for each of the top three quartiles of the "weight in kilograms/height in meters squared" distribution with the bottom quartile as the reference),
4. education (as one ordinal term),
5. marital status (as one binary term indicated whether the women had ever married),
6. age at joining the Adventist Church (as one ordinal term), and

7. index of participation in church activities (as one ordinal term).

Reproductive Factors

1. age at first live birth (using dummy terms for less than 25 years of age, 25 to 30, and 31 or older, with the group of women with no live births serving as the reference),
2. number of live births (as one continuous term),
3. combined number of miscarriages and stillbirths (as one continuous term),
4. age at menopause (using dummy terms for less than 40 years of age, 40 to 44, 45 to 49, and 55 and older, with the group of women with menopause at ages 50 to 54 serving as the reference).

Replacement Estrogens

1. replacement estrogen use (as a binary term indicating whether the subject had ever or never used replacement estrogens).

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APHA Job Placement Service

A Job Placement Service, facilitating the employment process for both applicants and recruiters, will be conducted during the 117th Annual Meeting of the American Public Health Association in Chicago, Oct. 22–26, 1989. Employers may list job vacancies and review resumes of candidates. Job applicants in turn, may submit resumes and review position descriptions. On-site interviews will be arranged at the employer's request. Those who are unable to attend the Annual Meeting, may purchase binders of available positions or applicant resumes. Registration forms will appear in the July and August 1989 issues of the Journal.

The Relation of Body Fat Distribution, as Assessed by Six Girth Measurements, to Diabetes Mellitus in Women

David S. Freedman, PhD, and Alfred A. Rimm, PhD

Abstract: Independently of the amount of adipose tissue, certain patterns of fat distribution increase the risk of non-insulin-dependent diabetes. Although the ratio of waist to hip (WHR) circumferences has been consistently related to diabetes mellitus, it is possible that only two measures do not completely characterize fat topography. The current study, therefore, examines the cross-sectional relation of six girths (waist, hip, neck, bust, wrist, and ankle) to diabetes mellitus in 43,595 women. As compared with non-diabetics, Quetelet index (kg/m^2) and all circumferences were elevated among diabetics. Stratified analyses showed that WHR, and waist, neck, and bust girths were consistently related to diabetes independently of the degree of overweight. As estimated from a logistic regression model that simultaneously controlled for age and all anthropometric variables, the prevalence of diabetes mellitus was positively related to Quetelet index, and to the waist, bust, and neck girths, with odds ratios (ORs) ranging from 1.4 to 2.6. However, diabetes was *inversely* related to hip (OR = 0.61) and ankle (OR = 0.73) girths; $p < 0.005$ for each association. Although cross-sectional in nature, these results suggest that an adverse body fat distribution is not limited to the abdominal region, but that a relative preponderance of adipose tissue in various regions of the upper body is associated with diabetes mellitus in women. (*Am J Public Health* 1989; 79:715–720.)

Introduction

Obesity is associated with numerous metabolic disorders and is often the primary cause of non-insulin-dependent diabetes.[1] Although obesity is a heterogenous entity,[2] systematic methods to characterize the distribution of body fat have long been in existence.[3] Vague, however, was the first to observe that, even at the same degree of overweight, women with a male (android) pattern of upper-body obesity have a greater prevalence of diabetes and atherosclerosis than do women with gynecoid obesity.[4] Android obesity was initially characterized by: 1) more subcutaneous fat at the nape of the neck than at the sacrum; and 2) a higher ratio of fat to muscle in the arm than in the thigh.[4,5]

However, other patterns of fat distribution have also been related to diabetes mellitus. Feldman, *et al*,[6] showed that female diabetics have a relative excess of subcutaneous fat in the subscapular, chest, and waist regions as compared to the triceps, and others[7–10] have confirmed the importance of a truncal fat distribution. Another body fat pattern, emphasizing abdominal obesity, has been defined by using the ratio of waist to hip (WHR)[11] or thigh[12] circumferences. An elevated WHR is related to the prevalence[13] and incidence[14] of diabetes mellitus, and to various metabolic complications.[15] Although additional validation studies are needed, both skinfold thickness and circumference measurements can be used to characterize the distribution of adipose tissue.[16]

Although other body fat patterns have been identified, several of these involve combinations of localized fat deposits, and are difficult to classify.[17–19] Because recent studies of diabetes mellitus have typically used only two skinfold or girth measurements to characterize body fat topography,[6,10,11,13–15] important aspects of fat patterning may have been overlooked. For example, WHR is only moderately correlated (r = 0.5 to 0.8) with Vague's android obesity.[5,20]

It is therefore possible that an adverse distribution of body fat is not restricted to the abdominal region (as assessed by WHR) but may encompass other areas in the upper body. The objective of the current study is to determine whether, in addition to the waist and hip girths, other circumference measurements are associated with diabetes mellitus in women independently of the degree of overweight.

Materials and Methods

Population

In 1969, approximately 125,000 questionnaires were distributed throughout the United States and Canada to local chapters of the Take Off Pounds Sensibly (TOPS) Club, an organization committed to weight reduction. Previously diagnosed diabetes was ascertained in the first part of the questionnaire, which was completed at a meeting of the local chapter. In the second part, completed at home, members were asked to list their current measurements for neck, bust, waist, hip, wrist, and ankle. (Forms were returned in sealed envelopes to assure confidentiality.) A detailed description of the study population, along with the questionnaire, has been published.[21,22]

A total of 73,532 White women completed the questionnaire, and the current analyses are restricted to non-pregnant, 20- to 70-year-olds (n = 67,902). The diabetic status of 3.7 per cent (n = 2,534) could not be determined, and data concerning one or more circumferences were missing for approximately one-third of the eligible women. In order to use the same set of women in all analyses, women with missing data for diabetes or any girth measurement were excluded. The sample size for all analyses was 43,595.

Of the women excluded because of missing girth measurements, wrist circumference was most likely to be incomplete (n = 21,277), whereas bust circumference was least likely to be missing (n = 9,059). Analyses were performed to determine whether these exclusions could have biased the results, and it was found that response was not related to any of the anthropometric measures nor to diabetes mellitus. For example, 4.7 per cent of the excluded women were diabetic, and their mean WHR was 0.04 higher than the mean WHR among non-diabetics. Comparable values were 4.5 per cent and 0.04 among women who were included in the study (see Table 1).

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TABLE 1—Comparison of Selected Characteristics between Nondiabetics and Diabetics

Characteristics	Percentiles:	Nondiabetics (n = 41,613)				Diabetics (n = 1,982)				t-Statistic[a]
		25	50	(mean)	75	25	50	(mean)	75	
Age (years)		31	39	(40)	48	38	46	(46)	54	21.7
Weight (kg)		66	74	(78)	86	73	85	(87)	98	21.3
Height (cm)		157	163	(163)	168	157	163	(163)	168	3.6
Quetelet Index (kg/m^2)		25.6	28.8	(30.0)	33.0	28.4	32.8	(33.8)	38.0	23.1
Circumferences										
Waist (cm)		74	81	(83)	89	81	91	(93)	102	29.2
WHR		0.72	0.76	(0.76)	0.80	0.76	0.80	(0.80)	0.84	24.8
Bust (cm)		94	99	(100)	107	97	107	(107)	114	24.6
Neck (cm)		33	36	(35)	36	33	36	(36)	38	21.6
Hip (cm)		102	107	(109)	114	104	114	(116)	124	19.4
Wrist (cm)		15	18	(17)	18	15	18	(17)	18	12.9
Ankle (cm)		20	23	(23)	25	23	23	(24)	25	8.0

[a]A t-statistic ≥ 3.3 indicates that the mean values are significantly different at the 0.001 level; a value ≥ 3.9 indicates statistical significance at the 0.00001 level.

Quetelet index (kg/m^2) was used as a measure of overweight.[23] Although weight was highly correlated with Quetelet index (r = 0.93), a cross-tabulation of quintiles of weight and Quetelet index indicated that Quetelet index (but not weight) was independently associated with the prevalence of diabetes. Furthermore, Quetelet index was strongly (r > 0.99) associated with relative weight, and was only weakly related (r = −0.08) to height. Mean values of relative weight across quintiles of Quetelet index were: 102, 114, 126, 140, and 172 per cent.

Disease Status

Diabetic status was determined from each subject's response to the question, "Has a doctor ever said that you have diabetes?"; 4.5 per cent (n = 1,982) responded "yes". Although these self-reported histories were not verified, agreement between self-reported gallbladder disease and physician's records has been examined: there was 13 per cent overreporting and 1.6 per cent underreporting by TOPS members.[24] Although the prevalence of undiagnosed diabetes in the US population may be substantial,[25] members of TOPS are likely to be more aware of chronic, obesity-related conditions than are other persons.

Women were also asked questions concerning the duration of diabetes and insulin usage. Of the 1,982 diabetics, 350 reported that they had been diagnosed within the last year and 322 reported using insulin. Because a large proportion of insulin users are non-insulin-dependent,[26] these latter women were included in all analyses. However, analyses restricted to non-insulin-using diabetics yielded similar results.

Statistical Analyses

Distributions of the anthropometric variables were non-normal. Neck, wrist, and ankle circumferences each had a high kurtosis, indicating a sharp peak with long tails; for example, 91 per cent of the women reported a wrist girth of six or seven inches. In contrast, distributions of Quetelet index and the truncal circumferences (bust, waist, and hip) were relatively flat. Both parametric and non-parametric tests were used in several analyses, but because of the large sample size, results were very similar.

Mean levels (and selected percentiles) of the anthropometric variables were first contrasted between non-diabetics and diabetics. Interrelationships among the circumferences were examined using Pearson correlation coefficients, and partial correlations were used to determine whether associations existed independently of the degree of overweight. Levels of Quetelet index and each girth measurement were then cross-classified, and independent increases in diabetes prevalence with each girth measure were assessed using the extended Mantel-Haenszel test for trend.[27]

Logistic regression was used to examine the relation of each girth measurement to diabetes, while simultaneously controlling for all other variables. To avoid assumptions concerning the shape of the associations, two indicator variables (representing contrasts among tertiles) were used for each predictor variable.[28] Estimated odds ratios (ORs) represent the odds of diabetes for women in the middle (or upper) tertile of each circumference vs the odds of diabetes for women in the lowest tertile.

Results

Mean levels and selected percentiles of age and the anthropometric variables are shown in Table 1. Diabetics were substantially older and heavier, with a mean Quetelet index that was 3.8 kg/m^2 higher than among non-diabetics. In addition, all circumference measures were larger among diabetics, with waist girth showing the largest mean difference (9.4 cm). In contrast, differences in ankle and wrist girths between the two groups were relatively small. Although the mean level of WHR among diabetes was only 0.04 higher than among non-diabetics, this ratio showed relatively little variability: the 25th and 75th percentiles among non-diabetics were 0.72 and 0.80, respectively.

Although diabetics had larger circumferences at all sites, assessing the independent relation of each girth measurement to diabetes was complicated by their strong interrelationships (Table 2). For example, Quetelet index and the waist, hip, and bust girths all showed correlations ≥0.80 with each other. In contrast, Quetelet index was only moderately associated with WHR (r = 0.28), and with wrist and ankle circumferences (r < 0.40). Except for WHR, age was only weakly related to the anthropometric variables. (Spearman correlations yielded very similar results; for example, the rank-order correlation between waist and hip girth was 0.82.)

Because all girth measurements increased with the degree of overweight, partial correlations were used to examine interrelationships among the girth measures independently of Quetelet index. Results in Table 2 (in parentheses) show that adjusting for overweight markedly reduced the intercorrelations among the girth measures. For example, correlations with hip girth decreased from r = 0.83 to 0.36 (for waist) and from r = 0.55 to 0.10 (for neck). Furthermore, wrist and ankle circumferences

TABLE 2—Interrelationships[a] among the Anthropometric Variables

	Quetelet Index	Age	Waist	Hip	WHR	Bust	Neck	Wrist
Quetelet index								
Age	.03							
Waist	.84	.14 (.22)[b]						
Hip	.88	.01 (−.03)	.83 (.36)					
WHR	.28	.25 (.25)	.62 (.74)	.09 (−.36)				
Bust	.81	.01 (−.03)	.85 (.53)	.80 (.31)	.40 (.30)			
Neck	.58	.09 (.08)	.62 (.30)	.55 (.10)	.34 (.22)	.59 (.25)		
Wrist	.36	.10 (.09)	.37 (.13)	.34 (.05)	.19 (.09)	.34 (.08)	.41 (.27)	
Ankle	.38	.06 (.05)	.35 (.05)	.39 (.11)	.09 (−.03)	.32 (0)	.36 (.19)	.45 (.36)

[a]Pearson correlation coefficients.
[b]Partial correlation coefficient, adjusted for Quetelet index and (Quetelet index)^2, is shown in parentheses.

were only weakly associated with the truncal (bust, waist, and hip) girths after controlling for Quetelet index. Additional adjustment for age did not alter these correlations.

The prevalence of diabetes was then examined after cross-classifying Quetelet index with each girth measurement (Table 3). As assessed by a test for trend (while controlling Quetelet index), diabetes was strongly related to WHR and waist, neck, and bust circumferences, with adjusted ORs ranging from 2.2 to 4.0 for women in the upper tertile of these girth measurements. Similar results were seen within each quintile of overweight: for example, the prevalence of diabetes increased from 1.9 to 6.0 per cent with increasing waist girths in the lowest Quetelet index quintile, and comparable increases were seen within other strata of Quetelet index. (Very little residual confounding was present in these analyses: as waist girth increased among women in the lowest overweight quintile, mean levels of Quetelet index were 23.4, 23.9, and 22.7 kg/m^2.) Although associations with WHR and both neck and bust girths were also similar across quintiles of overweight, diabetes was most strongly related to wrist and ankle girths among very overweight (quintiles 4 and 5) women. Among women with a Quetelet index >34.6 kg/m^2, the prevalence of diabetes *decreased* from 10.9 to 9.2 per cent with increasing ankle girth, and from 11.1 to 9.1 per cent with increasing hip girth.

Logistic regression was then used to determine which circumference measurements remained related to diabetes mellitus after controlling for all variables, and the final estimates of a stepwise procedure are shown in Table 4. Because dummy variables, representing tertiles of the pre-

TABLE 3—Prevalence of Diabetes Mellitus According to Levels of Quetelet Index and Girth Measurements

Girth Measurement	Tertile[b]	(N)	Quintiles[a] of Quetelet Index 1	2	3	4	5	Odds Ratios	χ^2 Test for Trend[c]
		(43,595)	2.1	2.5	3.5	5.5	9.1		
WHR	≤0.74	(15,210)	1.5[d]	1.7	1.9	3.1	4.7	1.0[e]	
	0.75–0.79	(14,188)	2.3	2.4	3.1	3.4	7.0	1.4	
	>0.79	(14,197)	4.3	4.0	3.5	8.4	12.2	2.8	302
Waist	≤76	(15,720)	1.9	1.9	1.8	2.1	0*	1.0	
	77–86	(13,745)	3.8	3.1	3.3	3.7	4.2	1.8	
	>86	(14,130)	6.0*	5.8	7.0	7.4	9.7	4.0	202
Neck	≤33	(21,338)	2.0	2.1	2.8	3.3	6.0	1.0	
	34	(13,163)	2.7	3.3	3.8	4.9	6.7	1.4	
	>34	(9,094)	4.3	4.0	5.4	5.8	10.9	2.2	144
Bust	≤94	(13,401)	1.9	2.3	2.2	3.5	3.6*	1.0	
	95–104	(17,359)	2.9	2.6	3.5	4.2	5.6	1.4	
	>104	(12,825)	4.2**	3.6	5.4	7.2	9.8	2.2	88
Wrist	<18	(21,067)	2.0	2.4	3.1	4.3	6.9	1.0	
	18	(19,663)	2.8	2.6	3.9	5.6	9.2	1.3	
	>18	(2,865)	1.2	2.2	3.3	10.3	11.5	1.8	41
Ankle	≤22	(11,348)	2.1	2.7	3.7	6.5	10.9	1.0	
	23	(20,179)	2.0	2.3	3.6	5.5	8.7	0.9	
	>23	(12,068)	2.6	2.5	3.2	4.9	9.2	0.9	4
Hip	≤102	(15,095)	2.1	2.6	4.0	7.9	11.1**	1.0	
	103–112	(14,773)	2.1	2.2	3.2	5.5	9.3	0.8	
	>112	(13,727)	3.4*	5.3	4.0	5.2	9.1	0.9	2

[a]Quintiles of Quetelet index (kg/m^2) are: < 25.1 (n = 8,689), 25.1–27.5 (n = 8,785), 27.6–30.3 (n = 8,675), 30.4–34.6 (n = 8,713), and > 34.6 (n = 8,733).
[b]All cut points for the circumference measurements are given in cm except for WHR.
[c]Test for trend.[28] With 1 degree of freedom, a value of 6.6 is statistically significant at the 0.01 level, and a value of 10.8 is statistically significant at the 0.001 level.
[d]Values represent percentage of women in each Quetelet index and girth category that are diabetic. All percentages are based on cell frequencies ≥ 100 except where noted: *< 100; **< 50.
[e]Reference category.

TABLE 4—Logistic Regression Analysis of Diabetes Mellitus and Girth Measurements

Predictor Variable[a]	Odds Ratio for Diabetes Relative to Women in Lowest Tertile[b]		Increase in −2 × Log likelihood[c]	P-Value
	Middle	High		
Waist	1.5	2.6	802	<0.0001
Age	1.8	2.8	270	<0.0001
Neck	1.1	1.7	99	<0.0001
Bust	1.1	1.6	35	<0.0001
Hip	0.69	0.61	21	<0.0001
Ankle	0.83	0.73	11	0.003
Quetelet index	1.2	1.4	12	0.002
Wrist	1.0	1.3	10	0.006

[a]Independent variables are listed in order of entry in a stepwise procedure.
[b]Each predictor variable was categorized into tertiles; cutpoints are shown in Table 3. Odds ratios are estimated from the final model which includes all predictor variables.
[c]−Log likelihood of model containing only the intercept is −8062 versus −7432 for full model. 2 × increase in −log likelihood for each predictor variable is approximately distributed as a chi-square statistic with 2 df.

dictor variables were used in these analyses, the estimated ORs indicate the odds of diabetes for women in the upper (or middle) tertile vs the odds of diabetes for women in the lowest tertile. Models that used interval scales for these variables yielded similar results (data not shown).

All circumference measurements provided additional information concerning diabetes. Although older (>47 years of age, upper tertile) women showed the largest OR (2.8) for diabetes, large waist girths were independently associated with an OR of 2.6. Large neck and bust girths further increased the risk of diabetes by 60 to 70 per cent, and a high Quetelet index by 40 per cent. However, at similar levels of the other variables, women with relatively large hip and ankle girths were *less* likely to be diabetic (ORs of 0.61 and 0.73, respectively). Additional models that included WHR (rather than waist and hip girths as two separate variables) yielded an OR of 1.9 for WHR; however, coefficients for the other variables were virtually identical to the estimates in Table 5.

The improved prediction of diabetes by the use of bust, neck, and wrist girths is shown in Figure 1 for the 35- to 46-year-olds (middle tertile of age). An OR of 3.9 was obtained if the prevalence of diabetes was contrasted between women in the upper and lower tertiles of Quetelet index, and this association increased to 7.3 if high levels of Quetelet index and WHR were contrasted with low levels of both variables. By using the bust, neck and wrist girths, the OR reached 9.7. However, in conjunction with these higher ORs (and positive predictive values), the sensitivity was reduced. Quetelet index alone had a sensitivity of 62 per cent (420/680) in identifying diabetic women, whereas the sensitivity of all five anthropometric variables was only 6 per cent (41/680).

Discussion

The importance of body fat distribution as an independent risk factor for diabetes mellitus has only recently been appreciated.[29,30] Although various skinfold and circumference measurements have been related to metabolic complications,[8–11,13–15] the current findings indicate that waist and hip girths do not completely summarize all of the detrimental aspects of adipose tissue distribution. After controlling for age and Quetelet index in the current study, diabetes mellitus was positively related to waist girth, and was inversely related to hip girth. However, even after accounting for these characteristics (or alternatively, for age, Quetelet index and WHR), diabetic women had larger neck, bust and wrist girths along with *smaller* ankle girths than did nondiabetics.

These findings indicate that diabetic women have a relative preponderance of adipose tissue in various regions of the upper (vs lower) body. Consistent with these results, Vague, *et al*, have argued that android obesity is not limited to one particular area, and correlations between android obesity and WHR are only moderate.[5,20] Furthermore, in

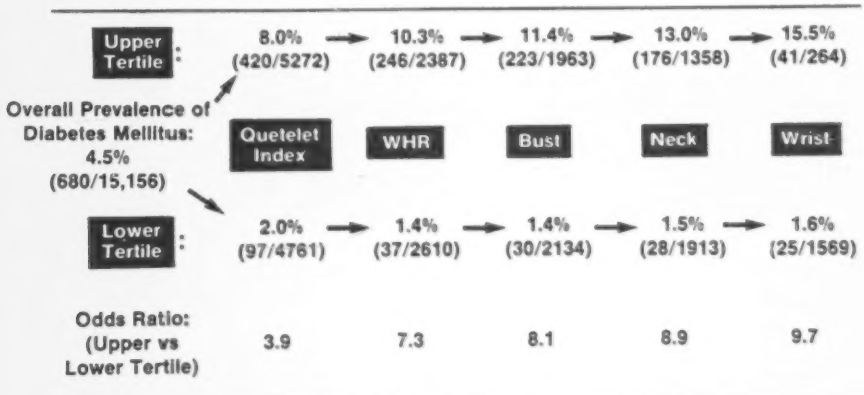

FIGURE 1—Improvement in the Prediction of Diabetes Mellitus among 35- to 45-year-old Women by the Use of Various Anthropometric Measurements

NOTE: The prevalence of diabetes for women in the upper vs lower tertile of each variable is contrasted. For WHR (waist to hip ratio) and subsequent variables, contrasts are restricted to women who were in the extreme tertiles for all previous measurements.

addition to WHR, other patterns of body fat distribution are deleterious: diabetics have relatively lean calf skinfolds[6,9] and an excess of subcutaneous fat in the subscapular region.[6,8,10] Because there are numerous patterns of fat topography that resist simple classification,[19] the current results suggest that a simple ratio cannot summarize all of the aspects of body fat distribution that increase the risk of diabetes mellitus.

Although girth measurements, particularly on the trunk, are related to the level of overweight, the importance of WHR (as an index of body fat distribution) has been demonstrated in studies that controlled for per cent body fat as determined from hydrostatic weighing.[31] Furthermore, after controlling for Quetelet index in the current study, both hip and ankle girths were *inversely* related to the prevalence of diabetes. Therefore, it is unlikely that circumference measurements merely provide a better indication of the total amount of adipose tissue. These findings, as do those of Mueller, *et al*,[16] indicate that combinations of circumference measurements are useful in describing body fat distribution. However, as compared with skinfold thicknesses, girth measurements are more reliable and can be assessed even in very obese persons.[32]

Although circumferences include other body components in addition to adipose tissue, individual truncal girths are strongly related to body fat. For example, Steinkamp, *et al*,[33] found that in middle-aged White women, body fat (as calculated from tritium dilution) was as strongly related to waist and chest girths as it was to the scapular skinfold thickness (r = 0.86 to 0.88). Furthermore, possibly because they include internal (as well as subcutaneous) adipose tissue, combinations of truncal circumferences are better predictors of total body fat than is ultrasound.[34] The observed relation of neck circumference to diabetes is in agreement with Vague's finding[5] of an increased skinfold thickness at the nape of the neck in women with android obesity.

In contrast, wrist and ankle girths are only weakly correlated (r = 0.1 to 0.6) with total body fat in women,[33,35] and are good measures of frame size in lean adults.[36] In these data, associations between diabetes and these circumferences were most evident among very overweight women (Table 3), suggesting that these peripheral girths may partly reflect regional fat deposition. The inverse relation of ankle circumference to diabetes is consistent with the contrast of upper- and lower-body obesity, and with the observation that diabetic women have relatively lean calves.[6,9] (Although ankle girths may be subject to greater misclassification than are other circumferences, nondifferential errors in measurement cannot explain its inverse association with diabetes mellitus.) However, it is possible that frame size, as assessed by wrist and ankle girths, is an additional risk factor: persons who are overweight without being obese may be at increased risk for hypertension and hypercholesterolemia.[37]

The association between both body fat patterning and clinical complications may be medicated by a third factor (sex hormones[15]), but there are several mechanisms whereby certain deposits of adipose tissue could result in metabolic disturbances. Decreases in both skeletal muscle sensitivity to insulin[38] and hepatic removal of insulin[15] are found in women with a high WHR, even after accounting for the degree of overweight. Furthermore, as compared with the femoral-gluteal region, abdominal fat cells show high rates of lipolysis.[39] Because free fatty acids from intra-abdominal adipocytes can drain directly into the portal circulation, the amount of intra-abdominal fat, which is highly correlated with WHR, may be the most important aspect of fat topography.[15]

Similar mechanisms may also be involved in other localizations of body fat; for example, mammary adipocytes also show high rates of lipolysis.[40] In agreement with the current findings concerning the importance of breast circumference in differentiating between diabetic and nondiabetics, previous studies have shown that: 1) the contrast of subscapular (rather than waist) and lateral calf skinfolds is optimal,[9] and 2) the subscapular skinfold thickness region is preferable to that in the abdominal region.[6] Furthermore, the subscapular skinfold is a slightly better predictor of coronary heart disease than is the abdominal skinfold.[41]

Although limitations of the cross-sectional study design should be considered, several of the possible biases would have reduced the magnitudes of the observed associations. Information was obtained by self-reports and it is likely that there was substantial misclassification of both diabetes and girth measurements. However, because women in TOPS were health-conscious, misclassification of diabetes was likely lower than that in the general population. Furthermore, any nondifferential misclassification would have resulted in an underestimation of the importance of girth measurements. Although it is also possible that overweight diabetics may have been more likely to join TOPS than were either overweight or diabetic women, this selection bias cannot explain the negative associations with hip and ankle girths, and because diabetic women may be motivated to lose weight, the observed associations may have underestimated the importance of body fat distribution. Additional analyses restricted to diabetics who were diagnosed within the last year (n = 461), however, yielded results that were similar to those observed among all diabetics.

Although few cohort studies have included baseline information on circumferences, the importance of WHR in the subsequent development of diabetes mellitus has been confirmed.[14] The current findings, however, indicate that an adverse localization of adipose tissue fat is not limited to the abdominal region, and that a generalized pattern of upper-body obesity is associated with diabetes mellitus in women. In addition to waist and hip circumferences, other girth measurements are needed to adequately characterize all of the detrimental aspects of body fat distribution.

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Optometry Journal Has New Name

The name of the journal of the American Academy of Optometry has been changed—from *American Journal of Optometry and Physiological Optics* to *Optometry and Vision Science*. Along with the shorter name, the journal changed its size and format as well. Now printed in the "A" size (8½ by 11 inches), the periodical bears the official color of the Academy—sea-foam green—and announced additional changes, including: larger, easier-to-read print, a new section dealing with statistics, more use of color, and more invited editorials.

Optometry and Vision Science is currently edited by William M. Lyle, OD, PhD, School of Optometry, University of Waterloo, Canada, and is published by Williams and Wilkins, 428 East Preston Street, Baltimore, MD 21202. Contact the publisher for a sample copy of the journal.

The Acceptance and Completion of Mammography by Older Black Women

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Abstract: We assessed the relation of patient characteristics, knowledge and beliefs to the utilization of mammography in an inner-city setting by 187 Black women over the age of 50. Thirty per cent of those who were offered mammography initially declined the offer and 40 per cent were subsequently unable to complete the procedure. Patient interviews were used to derive 27 potential knowledge and health belief predictor scales. In multiple regression analysis, two health belief scales and two knowledge scales accounted for 15 per cent of the observed variance in the model of acceptance. The strongest predictor of subsequent completion was initial acceptance. The presence of breast symptoms and two health belief scales together with initial acceptance accounted for 26 per cent of variance in the model of completion. These results suggest that the successful accomplishment of mammography requires coordinated efforts at the level of the provider, patient, and setting. Health beliefs may influence the patient's behavior in this process, but their effect appears to be modest. (*Am J Public Health* 1989; 79: 721–726.)

Introduction

Breast cancer is the leading cause of cancer mortality among Black women in the United States.[1,2] Programs of early detection can lower this risk by up to 40 per cent,[3,4] and thus standard medical practice protocols recommend periodic mammography for all women over the age of 50.[5,6] Nonetheless, only about 15 to 20 per cent of women over age 50 in the general population report ever having had a mammogram,[7–9] even though over 75 per cent of US women visit a physician each year.[10] No more than 4 per cent to 40 per cent of those studied in various medical settings receive mammography.[11–17] The accomplishment of mammography within these settings requires that physicians offer the procedure to patients who accept the recommendation within a framework which facilitates completion. Potential barriers include physician behavior (offering), patient behavior (acceptance), and logistic factors (such as cost, transportation, time).

We have previously noted that in a primary care setting serving an inner-city Black population, 20 per cent of female patients who were offered mammography initially declined the offer and nearly 40 per cent failed to complete the procedure.[18] Observations in our own and other settings have led to inconsistent conclusions concerning the relevance of patient health beliefs to participation in programs of early cancer detection.[18–23] We therefore conducted a prospective assessment of the relation of patient demographic characteristics, health status indicators, cancer control knowledge, and health beliefs to acceptance and subsequent completion of mammography among a group of older Black women in a primary care setting.

Methods

Subjects were recruited from patients visiting an Internal Medicine teaching practice during the period August through November 1985. The practice serves a population which is predominately Black, elderly, and of lower socioeconomic status. The practice was staffed by 16 resident physicians, one clinical nurse specialist, and three faculty internists, all of whom participated in the study.

At the conclusion of each patient visit during the study period the primary provider completed an encounter form which documented the provider's recommendation (offering) of any of three index procedures (Pap smear, mammography, home occult blood testing) to female patients over the age of 50. Following their visit, potential subjects were telephoned at home by a research assistant, informed consent for participation was requested, and those agreeing were then interviewed using the instrument described below. Of 474 potential subjects, 97 (20 per cent) declined participation, five (1 per cent) terminated the interview prematurely, and 72 (15 per cent) were unavailable (primarily due to absent phone service). The 300 participants were demographically similar to those declining the interview (mean age 67.4 versus 66.6 years) but did appear somewhat more likely to initially accept the offer of mammography (69 versus 63 per cent). Among the 300 interviewees, 207 had been offered mammography, and 276 were Black. The 187 Black interviewees offered mammography constitute the subjects of this report.

In our setting, mammography is provided via individually scheduled appointments. Following each visit during which mammography was recommended, the primary provider indicated the category of the patient's initial response as follows: *acceptance* (mammography appointment scheduled); *deferral* (patient request to delay consideration); *refusal* (patient rejection of the recommendation). Subsequent analyses contrasted those initially accepting the recommendation with those initially declining (defer plus refuse). Completion of mammography was documented through a review of the medical record and a computerized billing search conducted six months following the date of offering to allow an adequate interval for procedure completion.

The interview instrument employed was based upon the conceptual framework of the Health Belief Model[19] and included the dimensions of: knowledge, perceived severity, perceived susceptibility, perceived utility, perceived barriers, behavioral intentions, and cues to action. Items were derived from previous work by the authors and from review of previously published instruments, with particular emphasis on those assessing older and Black individuals.[18,20–28] The resulting final instrument contained 127 items and required no more than 45 minutes to complete.

Principal factor analysis with varimax rotation (SPSSx) was used to derive factors within each conceptual dimension (knowledge, susceptibility, etc). Internal consistency was determined using Cronbach's coefficient alpha which ranged

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from 0.45 to 0.97 with 24 of 27 factor alphas exceeding 0.60. Scales were then constructed by summing equally weighted item responses for each factor (see Appendix).

We measured the bivariate unadjusted relations of the demographic characteristics, health status indicators, interview scales, to mammography utilization via Pearson correlation and stratified analyses contrasting the effect of scale values classified as high, moderate, or low (generally employing cutpoints at the 75th and 25th percentiles, depending upon the frequency distributions). In preliminary regression analyses, we fitted models for each individual block of variables (demographics, health status indicators, knowledge and health belief domains) with acceptance or completion as the dependent variable. We then developed separate individual models of acceptance and completion, employing only those predictor variables which seemed reasonable and were independently significant at the 0.10 level in the initial analyses (Table 2). Results from linear models are presented here; logistic analyses gave similar results. Estimates of marginal and composite probability were derived from the logistic models.[29] The SPSSx and BMDP software packages were used for all statistical analyses and two-tailed tests of significance are reported.

Results

The average age of the 187 interviewees was 67.1 years (S.D. = 10.6). Seventeen per cent were currently married and half lived alone. Nearly 70 per cent reported household incomes less than $5000, 11 per cent were currently employed, the mean education level was 9.0 years (S.D. = 3.1), and all had some form of third party payment. The prevalence of chronic illness was high: over three-fourths had hypertension, nearly one-half manifest heart disease, one-third had diabetes mellitus, and 40 per cent had osteoarthritis. All were ambulatory and residing in the community.

Of the 187 subjects offered mammography, 131 (70 per cent) initially accepted, 42 (22 per cent) deferred, and 14 (7 per cent) refused the offer. Overall, 109 (60 per cent) of patients offered mammography were able to complete the procedure during the six-month period of follow-up. Completion was more likely among those initially accepting the offer, but some who initially accepted did not complete mammography and vice versa. Overall, 75 per cent of those initially accepting the offer completed mammography, as contrasted with 29 per cent of those initially deferring, and 8 per cent of those initially refusing.

Results from the patient interviews are presented in the Appendix. Overall, knowledge of cancer control was limited although 90 per cent believed that early detection in general was useful. Only one-third identified breast cancer as curable if detected early. Thirty per cent mentioned mammography as a potentially useful early detection procedure and half could accurately define mammography.

Most women perceived the lifetime risk of breast cancer in the population as two to four times the actual risk (median response 25–50 per cent), yet only 5 per cent viewed cancer as very likely in their own lifetime and only 1 per cent specifically identified cancer of the breast as a future health concern. Still, nearly 50 per cent acknowledged that they worried about getting cancer or that tests might detect cancer. Furthermore, 60 per cent felt they had no control over getting cancer and that there was "no point in worrying about cancer now" since it was unavoidable.

Perceived barriers to early detection included the expense of an office visit (40 per cent) or of mammography (23 per cent). Mammography safety, discomfort, or fear were each cited by less than 13 per cent as having any importance as potential barriers.

The interrelations among the potential predictor variables were relatively weak for the most part, thus limiting the potential effect of multicollinearity. Age, education, and knowledge scores did demonstrate relationships with selected health beliefs in the expected direction. Overall, of the 406 interscale bivariate correlation coefficients only 31 exceeded 0.3 and, of these, three exceeded 0.6, i.e., *utility of tests* (likelihood of behavior change); *emotional arousal* (level of concern); *futility of cancer treatment* (knowledge of prognosis).

The unadjusted relation of selected demographic and interview scale variables to the acceptance and completion of offered mammography is presented in Table 1. Older, asymptomatic women with more frequent medical conditions were less likely to accept and complete mammography. Acceptance and completion were also associated with: knowledge of the procedure; belief that testing, early detection, and treatment were useful; and belief that cancer was avoidable. Women who discounted the importance of expense, fear of detection, embarrassment, and procedure discomfort as barriers were more likely to accept and complete mammography.

The results of the multiple regression analyses are presented in Table 2. The previously described bivariate intercorrelations noted between predictor variables and the relatively skewed distribution of some of the belief scores (Table 1, particularly barriers 2–5) accounted for the exclusion of the variables in Table 1 that were not in the final multivariate models. We observed no independent relation between any of the demographic or health status indicators with initial acceptance of mammography. Women who knew more about the detection procedures and recognized the increased risk of cancer with age were more likely to accept, other factors being equal. Health beliefs independently associated with acceptance included the belief that early treatment of cancer was useful, and a discounting of the importance of procedure expense as a potential barrier to mammography accomplishment. Overall, these four scales accounted for 14.5 per cent of the observed variance in mammography acceptance (95% CI = 7.7, 34.4).

Initial acceptance strongly predicted mammography completion, as expected. One health status indicator and two health belief scales also contributed independently to the final model. The presence of breast symptoms, discounting of expense as a barrier, and higher assigned utility of cancer detection tests, together with initial acceptance, accounted for 26 per cent of the observed variance in the final model of mammography completion (95% CI = 18.8, 55.3).

As shown in Table 3, the magnitude of the independent effects (marginal probabilities) of each of the four predictor variables in the acceptance model varied between 5 per cent and 9 per cent. For example, a one-point increase in the health beliefs scale for utility of treatment was linked to an increase of 8 percentage points (95% CI = 2.9, 13.1) in acceptance. The presence of symptoms was linked to an increase in completion of 15.5 percentage points (90% CI = .6, 30.4). Table 3 presents estimates of the probability of acceptance and completion, based upon the logistic models, for various combinations of predictor variables. The estimated probability of mammography acceptance among those with higher belief scores rose from 66 per cent among those with low scores on the knowledge scales to 92 per cent for

TABLE 1—Percentage of Women Accepting and Completing Recommended Mammography, According to Selected* Interview Scores and Other Factors

Factors	Mammography Offered (n)	Mammography Accepted (%)	Mammography Completed (%)
All subjects	187	70.1	59.6
Age (years)			
50–59	55	79.2	58.2
60–69	60	66.7	70.7
70–79	44	72.7	59.5
≥80	28	57.1	39.3
Income			
≤$5000	130	76.2	62.2
>$5000	51	58.8	56.0
Serious Illnesses (n)			
0	100	73.8	59.0
1	69	71.0	65.2
≥2	14	40.0	35.7
Breast Symptoms			
Present	28	82.1	78.6
Absent	159	67.9	56.1
Knowledge 2—Procedure			
Low (3.0–4.0)	60	50.0	68.2
Moderate (4.1–6.2)	70	57.1	62.5
High (6.3–7.2)	48	72.9	77.9
Knowledge 5—Age Risk			
Low (2.0–2.9)	85	60.0	55.4
Moderate (3.0–3.9)	23	87.0	73.9
High (≥4.0)	79	75.9	59.7
Susceptibility 3—Unavoidable			
Low (4.0–6.9)	46	71.7	66.7
Moderate (7.0–10.9)	81	77.8	66.3
High (≥11.0)	54	57.4	42.3
Barriers 1—Expense			
Low (3.0–5.9)	135	75.6	67.7
Moderate (6.0–8.9)	21	66.7	42.9
High (≤9)	23	43.5	39.1
Barriers 2—Fear			
Low (≥9)	155	73.5	62.7
High (3.0–8.9)	25	52.0	52.0
Barriers 3—Embarrassment			
Low (≥9)	141	66.0	73.2
High (3.0–8.9)	37	43.2	56.7
Barriers 5—Discomfort			
Low (≥9)	146	76.0	64.6
High (3.0–8.9)	29	44.8	44.8
Utility 1—Tests			
Low (4.0–11.9)	22	40.9	22.7
High (≥12)	164	74.4	64.6
Utility 2—Early Treatment			
Low (4.0–10.9)	23	56.5	47.8
Moderate (11.0–11.9)	31	58.1	53.3
High (≥12)	127	76.4	63.7
Utility 3—Futility Treatment			
Low (4.0–6.9)	40	87.5	75.0
Moderate (7.0–10.9)	94	70.2	58.9
High (11.0–12.9)	42	47.6	45.2
Utility 4—Futility Detection			
Low (4.0–8.9)	68	80.9	67.2
Moderate (9.0–11.9)	65	64.6	58.1
High (≥12)	47	59.6	48.9

*Significant at the 0.10 level in bivariate analyses.

those with high scores. Table 3 also includes the observed rates of acceptance and completion for comparison with the model estimates. It should be noted that the small number of individuals manifesting concurrently low values on each of the relevant health belief scales precludes calculation of rates for those groups.

Discussion

Successful cancer control in the primary care setting requires that detection procedures be offered by physicians to patients who then accept and are able to complete the procedure. We have observed that even among women offered the procedure, personal and logistic factors are important determinants of subsequent acceptance and completion of mammography. While provision of mammographic services to a substantial number of the patients we studied is encouraging, those remaining at risk for nonparticipation constitute an important focus for study.

The role of health beliefs in this process is complex. Our data suggest that patient beliefs only weakly influence initial acceptance and subsequent procedure completion. After controlling for the interdependence of the multiple potential predictors, only two knowledge variables (detection procedures, age risk) and two belief scales (utility of early treatment, importance of expense as a barrier) remained as independent predictors of mammography acceptance and together accounted for 15 per cent of the observed variance. After controlling for initial acceptance, subsequent completion appeared more likely among women who presented with a sign or symptom of breast disease, who believed in the utility of early detection tests and who discounted expense as a barrier. The inverse relation noted for external prompts may suggest that intrapersonal factors are the more relevant but this must be interpreted with caution. While this model explained 26 per cent of the variance in completion, initial acceptance accounted for the largest component. Previous studies of patient participation in early detection programs for breast as well as colon cancer have also reported only modest effects for health beliefs.[22,23] These observations lead us to conclude that factors other than patient beliefs must be considered if mammography utilization is to improve our setting.

Mammography appears substantially underutilized as a cancer control tool in most reported settings.[11–17] Our findings emphasize that the process of mammography utilization begins with a physician decision to recommend the procedure but then depends upon the successful interaction of patient, physician, and setting. McCusker and Celantano have suggested that physicians may operate as potent reinforcers of patient decisions regarding early cancer detection, perhaps consistent with our observation of subsequent examination completion among women initially deferring or refusing the procedure.[21,30] We have not, however, assessed the interactional dynamics of the offers, nor those factors associated with the physician's decision to recommend mammography to an individual patient.

This study and previous observations have also emphasized the importance of logistic factors such as procedure cost as potential barriers.[31] Although all of our subjects had health care insurance, coverage for preventive services is variable and we do not have information on individual out-of-pocket expense. Nor have we directly ascertained the extent to which other logistic factors inhibited patient intention to complete mammography. Retrospective interviews may be useful to assess experienced barriers and would complement prospective ascertainment of predictive characteristics.

Our focus has been the relevance of potential barriers to the utilization of mammography among older, Black women and thus our conclusions may not apply to other settings. Education, knowledge of cancer control, and resources were limited among our subjects. National surveys, however, have also demonstrated a surprising degree of misinformation regarding cancer among many groups, including but not limited to older and minority subgroups.[24,28]

TABLE 2—Results of Multiple Linear Regression Analyses for the Acceptance and Completion of Mammography

	Mammography Acceptance			Mammography Completion		
	b	(90% CI)	Marginal Probability*	b	90% CI	Marginal Probability*
Presence of Symptoms	—	—	—	.155	(.006, .304)	.120
Health Beliefs						
Utility treatment	.080	(.038, .123)	.087	—	—	—
Utility tests	—	—	—	.042	(.008, .076)	.068
Expense barrier	.044	(.016, .071)	.049	.031	(.003, .059)	.040
External prompts	—	—	—	−.072	(−.146, .002)	—
Knowledge						
Procedure	.073	(.031, .115)	.084	—	—	—
Age risk	.062	(.002, .121)	.080	—	—	—
Acceptance	—	—	—	.437	(.312, .562)	.256
Constant	−1.120	(−1.702, −.534)	—	−.562	(−1.033, −.092)	
Multiple R	.407			.529		
Adjusted R^2	.145			.257		

*Based upon logistic regression model
NOTE: N = 166

TABLE 3—Estimated and Observed Probabilities of the Acceptance and Completion of Mammography

	Mammography Acceptance*				Mammography Completion**			
	High Knowledge		Other Knowledge		Initial Accept		Initial Decline	
Health Belief Scales***	Estimate %	Observed % (n)	Estimate %	Observed % (n)	Estimate %	Observed % (n)	Estimate %	Observed % (n)
High	92	94 (34)	67	75 (56)	89	82 (79)	63	22 (23)
Other	66	69 (26)	43	53 (53)	57	41 (22)	32	28 (25)

*Overall probability of acceptance = .71
**Overall probability of completion among asymptomatic women = .58
***Scales in acceptance model include utility of treatment and discounting expense barrier. Scales in completion include utility of tests and discounting expense barrier.

Several other limitations of this study must be acknowledged. Nearly one-third of all potential subjects were not interviewed. While the initial mammography acceptance rate among those not interviewed was similar to that observed among the participants, we have no information concerning their subsequent behavior or their beliefs. Furthermore, the response rate specifically among Black women could not be measured. Second, health beliefs are complex constructs and no reference standard for their assessment is available. In addition, incorporation in our models of unmeasured but relevant variables or study of a sample manifesting broader variability could alter our estimates. Finally, even if these results are correct for this population, we do not know how generalizable they are.

Successful cancer control in the primary care setting will likely require complementary interventions at the level of patient, provider, and institution. Structural interventions which limit cost or which otherwise facilitate completion are fundamental. Certainly, patient beliefs are not irrelevant to this process. Furthermore, while their influence may be limited, health beliefs are potentially modifiable and thus of clinical relevance particularly in a population characterized by limited knowledge of cancer control activities. However, our work and that of others reinforces the need to interpret patient beliefs in the broader context within which decisions are made and action carried out. This broader view of physician and patient behavior holds the most likely promise for cancer control in the primary care setting.

ACKNOWLEDGMENTS

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APPENDIX
Description of the Patient Interview Scales

Health Belief Factors	Description	Mean	Range
Knowledge			
1. Recognition of symptoms (5 items, alpha = .78)	Breast and colon symptoms suggestive of cancer	8.89	5–10
2. Procedures (4 items, alpha = .72)	Identification and purpose of Pap smear, mammography, and occult blood testing	5.11	3–8
3. Routine checkup (2 items, apha = .89)	Indication for routine examinations	2.98	2–4
4. Prognosis (3 items, alpha = .64)	Potentially curable cancers and likelihood of fatality	4.96	2–9
5. Age effect (2 items, alpha = .86)	Risk of cancer with increasing age	2.94	2–4
6. Prevalence (2 items, alpha = .66)	Population likelihood of breast and colon cancer	.21	0–2
Severity			
1. Limitations (3 items, alpha = .78)	Impact of cancer on social activity	6.24	3–9
2. Emotional arousal (3 items, alpha = .66)	Fear of cancer severity	6.77	3–9
Susceptibility			
1. Comparative risk (4 items, alpha = .90)	Likelihood of personal cancer compared to others	7.54	4–12
2. Absolute risk (2 items, alpha = .91)	Personal likelihood of cancer over time	6.22	3–9
3. Unavoidability (4 items, alpha = .67)	Potential to prevent or avoid cancer	8.22	4–12
4. Level of concern (5 items, alpha = .55)	Worry regarding personal likelihood of cancer and illness	5.73	3–11 25
Utility			
1. Worthwhile tests (4 items, alpha = .97)	Usefulness of tests	11.47	4–12
2. Early treatment (4 items, alpha = .72)	Potential for cure if cancer detected early	11.41	4–12
3. Futility cancer (4 items, alpha = .70)	Futility of cancer treatment	8.30	4–12
4. Futility detection (4 items, alpha = .56)	Futility of early cancer detection	9.99	4–12
Barriers			
1. Procedure expense (3 items, alpha = .95)	Importance of cost in detering utilization	7.93	3–9
2. Fear (3 items, alpha = .89)	Importance of fear of detection in detering utilization	8.63	3–9
3. Embarrassment (3 items, alpha = .84)	Importance of embarrassment in detering utilization	8.53	3–9
4. Safety (3 items, alpha = .86)	Importance of procedure safety in detering utilization	8.45	3–9

(continued)

APPENDIX (continued)

Health Belief Factors	Description	Mean	Range
5. Discomfort (3 items, alpha = .74)	Importance of discomfort in detering utilization	8.62	3–9
6. System devaluation (3 items, alpha = .67)	Discounting of physician or system	8.64	3–9
7. Resource limitations (3 items, apha = .66)	Relevance of expense and transportation as obstacles	5.57	3–9
8. Time (3 items, alpha = .45)	Relevance of time as an obstacle	8.56	4–9
Behavioral Intentions			
1. Likelihood change (3 items, alpha = .81)	Probability of altering diet or habits to reduce cancer risk	8.31	3–9
2. Willingness checkup (2 items, alpha = .60)	Likelihood of engaging in regular checkups	5.66	2–6
Cues to Action			
1. Physician recommendation (3 items, apha = .73)	Recollection that personal physician endorses procedure	7.50	3–9
2. Cancer experience (3 items)	Personal contact with individual with cancer	1.19	0–3
3. External prompts (3 items)	Suggestions by individuals or media regarding checkups	.71	0–3

CDC's ACIP Updates General Recommendations on Immunization

The Immunization Practices Advisory Committee (ACIP), of the Centers for Disease Control, has updated its general recommendations on immunization. The revised version, which updates the 1983 statement, was printed in the April 7, 1989 issue of *Morbidity and Mortality Weekly Report* (MMWR) published by CDC. Changes or new sections in the revised recommendations include:

1) Listing of vaccines available in the United States by type and recommended routes of administration;
2) Updated schedules for immunizing infants and children;
3) Clarification of the guidelines for spacing administration of immune globulin preparations and different vaccines;
4) An updated table of recommendations for routing immunization of children infected with human immunodeficiency virus;
5) Listing of conditions that are often inappropriately considered contraindications to immunization; and
6) Addition of information on the National Childhood Vaccine Injury Act of 1986 and the National Vaccine Injury Compensation Program.

The *Morbidity and Mortality Weekly Report* is prepared by the Centers for Disease Control, Atlanta, Georgia, and available on a paid subscription basis from the Superintendent of Documents, US Government Printing Office, Washington, DC 20402, (202) 783-3238.

Psychiatric Disorders and 15-Month Mortality in a Community Sample of Older Adults

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Abstract: The relation between psychiatric disorders, defined by the Diagnostic Interview Schedule (DIS), and mortality over 15 months is compared in 3007 adults age 55 and over in the New Haven Epidemiologic Catchment Area (ECA) project. Our results indicate that the odds of dying are more than four times greater for individuals with affective disorders than for others in the sample, controlling for age, sex, and physical health. Cause of death is also examined. There were no suicides or deaths from external causes. (*Am J Public Health* 1989; 79:727–730.)

Introduction

In the past decade, a number of studies have examined the prevalence of psychiatric disorders in the community.[1,2] While these disorders exact their toll on the quality of individual lives, follow-up studies of patients in treatment suggest that certain psychiatric disorders also increase the risk for death.[3–5] However, only a few studies have examined the effects of specific psychiatric disorders on mortality in community samples and these yield inconsistent results.[6,7]

The Epidemiologic Catchment Area (ECA) project was designed to assess the prevalence of major psychiatric disorders in community samples. In its follow-up, the study collected mortality data, providing a unique opportunity to examine the relationship between overall and cause-specific mortality and psychiatric disorders in the older members of the sample while controlling for a number of known predictors with mortality.

Methods

Data for these analyses were collected as part of three waves of the New Haven portion of the ECA project, a collaborative multi-wave study of the incidence and prevalence of major psychiatric disorders and the use of health and mental health services. Beginning in July 1980, interviews were obtained from a multistage probability sample of 3,058 adults age 18 years and over living in a 13-town region comprising the greater New Haven community (total adult population of approximately 300,000). An additional sample of 1,976 elderly respondents age 65 and over was also assessed. The sampling frame stratified data collection among the 13 towns; household and elderly oversample respondents were selected from among clusters of eight households at intervals of 60 households within each of the 13 strata. Only one respondent was interviewed from a household. The methods used in ECA project are described in greater detail elsewhere.[8–10] The total of 5,034 interviews represents an overall response rate for the first wave of the survey of 77 per cent. Because the observed 15-month mortality is so low in the younger (ages 18–54) age group (three deceased or 0.1 per cent), we examine mortality only among the 3,007 respondents age 55 and over.

Two follow-up contacts were scheduled in an average follow-up period of 15 months. In this analysis, wave 1 respondents identified by informants as having died over the course of the follow-up period are compared to the remaining wave 1 sample. This approach assumes that respondents lost to follow-up survived the 15-month period. However, because the current study is preliminary to a planned long-term follow-up of the full wave 1 sample, we have not been able to confirm this assumption. The group lost to follow-up includes: 48 respondents not found at the second wave; 497 respondents who were alive at the second wave but refused to be reinterviewed and were therefore not contacted at the third wave; and 14 respondents who were interviewed at the second wave but could not be located at the third wave. These 559 respondents did not differ from the rest of the sample by age, sex, or lifetime or recent psychiatric disorders although they were more likely to be in good physical health or to be cognitively impaired. Including them in these analyses as living respondents results in a more conservative estimates of the effects of psychiatric status on mortality than omitting them from the analysis. With this group omitted, the only difference in results was an increased effect of recent schizophrenia on mortality in the comorbidity model.

During this period, 136 (4.5 per cent) of our sample of adults age 55 and older were identified by informants as having died. This is 25 fewer deaths than the number expected based on national age-sex-race specific mortality rates in 1980, suggesting that there may have been some deaths among the respondents lost to follow-up. Using local and state resources, 131 of the 136 (97 per cent) of the deaths have been verified by death certificates. Cause of death was extracted from the death certificates and classified with the assistance of a physician. Of the unconfirmed deaths, one reportedly occurred outside of the United States and at least one occurred to respondents who moved from Connecticut during the study period.

Psychiatric status was assessed by the Diagnostic Interview Schedule (DIS), a structured interview designed for use by lay interviewers in epidemiologic studies and capable of generating diagnoses for the more prevalent psychiatric disorders.[11] The DIS ascertains information about both the existence and severity of symptoms, and computer algorithms determine whether these meet DSM-III* criteria for specific psychiatric disorders. Severe impairment in cognitive functioning at the time of the interview was assessed by the Mini-mental State Examination.[12] The likelihood of mortality is compared using logistic regression for individuals

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*Diagnostic and Statistical Manual of Mental Disorders, American Psychiatric Association.

meeting lifetime or recent (within six months of the interview) criteria for any of four groups of psychiatric disorder or for cognitive impairment. The disorder groups examined include affective disorders, anxiety disorders, substance abuse or dependence, and schizophrenic disorders.

Controls include age (in years), sex, and baseline physical health status. Physical health status was assessed by asking respondents whether they had a history of any of a variety of chronic health problems (hypertension, heart attack, stroke, cancer) or any limitation in physical functioning.[13] Approximately 80 per cent of the sample of adults age 55 and over reported physical health problems using these criteria. Although this represents a gross assessment of health, the crude relative risk of 15-month mortality for respondents age 55 and over associated with poor health is 6.38, a figure comparable or greater than observed with measures used in other studies of older adults.[14,15]

Results

As seen in the first column of Table 1, mortality rates varied considerably by psychiatric status. The percentage deceased was greatest among those with affective disorders (10.9 per cent for those meeting criteria within six months of the interview), schizophrenic disorders (14.3 per cent for recent and 6.3 per cent for lifetime), and severe cognitive impairment (9.0 per cent). Adults age 55 and over who had an affective disorder, schizophrenia, or cognitive impairment within six months of the initial interview had at least 150 per cent greater chance of dying in the following 15 months than did other members of the sample. The odds ratios also indicate the relatively low risk of mortality in this sample among those with anxiety disorders (0.9 for recent and lifetime) or who abuse drugs or alcohol (0.8 for recent and 1.1 for lifetime).

The far right column in Table 1 presents the odds ratio and 95% confidence interval of dying (adjusted for sex, age, and physical health status using logistic regression). These adjusted estimates indicate that the high mortality associated with affective disorders (2.97 for recent and 1.83 for lifetime) was independent of sex, age, and physical health status. The association with schizophrenia is also strong (3.94 for recent and 1.82 for lifetime disorders) but the small number of cases (N = 7) yielded an extremely wide confidence interval. In contrast, the risk of dying associated with cognitive impairment is reduced considerably, from 2.5 to 1.43, once sex, age, and physical health status are controlled.

The effects of comorbidity among psychiatric disorders were controlled using a logistic regression model with all recent psychiatric disorder groups and cognitive impairment entered with sex, age, and physical health status. As seen in Table 2, men were 2.83 times more likely to die than women over the course of the 15 months, the likelihood of dying increased with age, and those in poor physical health are 4.31 times more likely to die than those in good health. The odds ratios associated with each of the disorder groups were similar to the adjusted estimates reported in Table 1 except for the affective disorders which increased to 4.31 when controlling for comorbidity.

To test whether the effects of recent affective disorders on mortality were stronger when accompanied by physical health problems,[15] we added an interaction term to the logistic regression model reported in Table 2 even though there were few affective cases in each physical health status cell. The interaction was significant ($p < .01$), although not in the expected direction, as illustrated by the crude mortality rates disaggregated by affective and physical health status in Table 3. Affective disorders increased the odds of mortality for those both with and without physical health problems, but the effect was greater for those reporting no physical health problems.

When we stratified the recent affective cases by type of disorder, we found that all the deaths were among those individuals meeting criteria for major depressive disorder (without a history of manic episodes) and none were among those with bipolar disorder. Of recent major depressives, five (12.5 per cent) died during the study period. Among lifetime cases not reporting a recent episode, one of the 36 (2.8 per cent) with major depression and one of the three (33 per cent)

TABLE 1—15-Month Mortality by Psychiatric Status and Cognitive Functioning

	N	% Deceased	Odds[a]	Adjusted Odds[b] (95% CI)
DIS/DSM III Disorder Groups				
Affective				
Lifetime[c]	134	6.7	1.6	1.83 (0.78–3.29)
Recent	46	10.9	2.7	2.97 (1.12–7.92)
Anxiety				
Lifetime	208	3.4	0.9	1.07 (0.49–2.37)
Recent	144	3.5	0.9	1.02 (0.40–2.52)
Substance Abuse				
Lifetime	126	4.8	1.1	0.99 (0.42–2.37)
Recent	28	3.6	0.8	0.81 (0.11–6.26)
Schizophrenia/form				
Lifetime	16	6.3	1.4	1.82 (0.22–15.09)
Recent	7	14.3	3.5	3.94 (0.37–41.36)
Severe Cognitive Impairment	111	9.0	2.5	1.43 (0.70–2.95)
No DIS/DSM III Disorder or Cognitive Impairment	2,479	4.3	0.7	0.79 (0.52–1.22)
Age	3,007	[d]	[d]	4.06[e] (2.63–6.24)
Male	1,196	6.4	2.0	2.60 (1.82–3.73)
Physical Health Problems	2,419	5.4	6.7	5.17 (2.08–12.77)

[a]Odds compare index group vs all others.
[b]Odds adjusted for sex, age (in years), and physical health problems using logistic regression.
[c]Lifetime and recent are overlapping categories.
[d]Mean age of deceased is 76.6 years; mean age of nondeceased is 71.2 years.
[e]Odds Ratio compares risk of 20-year difference (e.g. 75 years to 55 years).

TABLE 2—Factors Predicting Mortality in a 15-Month Period, Logistic Regression Model

	Logistic Odds	Odds Ratio	95% CI
Male	1.04	2.83	1.87, 4.27
Age (years)	0.07	4.06[a]	2.74, 6.00
Physical Health Problems	1.46	4.31	1.41, 13.16
Recent DIS/DSM-III Disorders			
Affective	1.46	4.31	1.71, 10.82
Anxiety	−0.20	0.82	0.31, 2.14
Substance	−0.12	0.89	0.12, 6.81
Schizophrenia/form	1.56	4.76	0.46, 49.00
Cognitive Impairment	−0.29	0.75	0.26, 2.18

[a]Odds Ratio compares risk of 20-year difference (e.g., 75 years old to 55 years old).
NOTE: Model intercept = −8.46, N = 2875.

with bipolar disorder had died. Few adults over age 55 in our sample had a history of bipolar disorder but, in these data, the risk of mortality among those with affective disorders was strongest for those with major depression, especially recent episodes.

Finally, we obtained cause of death from death certificates and compared the causes of death between psychiatric cases and noncases. Among the deceased, 63 per cent had died from vascular-related (cardio and cerebral) problems, 22 per cent from cancer, and the remaining 15 per cent from other causes, primarily pulmonary diseases. We found no evidence that any of the respondents committed suicide or died from other unnatural causes. Overall, the distribution of causes of death did not differ between psychiatric cases and noncases. While there was some variation in the distributions of causes of death among the psychiatric disorder groups, the small number of cases in each precludes generalizing from these observations. Among the depressed respondents who died during the study period, two died from heart-related problems, two from cancer, and one from respiratory arrest due to septic shock.

Discussion

These data suggest that premature mortality is associated with some, but not all, psychiatric disorders for adults age 55 and over. In this study, the risk was greatest for individuals meeting criteria for an affective disorder, specifically major depressive disorder, within six months of their interview. Because these analyses controlled for physical health status, the increased odds of death for individuals with an affective disorder cannot be attributed solely to comorbidity with physical health problems. Moreover, although the number of individuals with an affective disorder who report no history of physical health problems was small, the risk of mortality was at least as great for this group as for affective cases reporting physical health problems.

TABLE 3—Percentage Deceased (unadjusted) by Physical Health Status and Affective Disorders

	Recent Affective Disorder			
Physical health problems	No	%	Yes	%
No	(4/581)	0.7	(1/4)	25.0
Yes	(126/2368)	5.3	(4/42)	9.5

NOTE: Interaction between physical health and affective disorders significant ($p < .01$) adjusting for sex and age.

This examination of the effects of DSM-III diagnoses on mortality in a recent community sample of both treated and untreated persons augments the findings derived from groups of psychiatric patients,[4,5] from community samples which assessed depressive symptoms rather than diagnoses,[16] or from community samples in earlier decades.[6] In a review by Tsuang and Simpson,[3] Standardized Mortality Ratios in several patient studies of affective disorders range from less than 1 to 2.5 and in the follow-up of the 1952 Stirling County sample are approximately 1.6;[6] the relatively larger Odds Ratios of 4.31 for affective disorders found in the ECA data contradict the notion that the risk associated with these disorders for mortality is decreasing with time.

These findings are disturbing given the major advances in the treatment of affective disorders over the past three decades. Some evidence indicates that treatment can reduce the risk of premature death in affective patients.[17] Yet when we examined the treatment history of the affective cases who died in the 15 months following their initial interview, none reported being treated for mental health problems by a mental health specialist or general medical practitioner in the six-month period prior to the first interview and only one reported such treatment by a medical practitioner after the interview. While low usage of mental health services is consistent with older age groups in general,[9,10] the high risk posed by affective disorders warrants increased attention of all medical practitioners and care-givers to identifying and treating mental and emotional problems in older adults.

While this study confirms the risk associated with affective disorders for mortality even in recent cohorts of older adults, the data tell us little about the process by which this risk is increased. For example, for some individuals not evidencing physical problems, the depressive symptoms may be prodromal to conditions which will manifest in future months. For others, the impairment associated with depression may result in poor hygiene and nutritional neglect, increasing a person's susceptibility to physical health problems or exacerbating existing health problems. A longer period of follow-up will be useful in determining whether the high risk observed in this relatively short period persists over time.

In contrast, the apparent effect of cognitive impairment on mortality was largely explained by other factors. Because this result contradicts other studies, we examined whether the lack of effect reflected comorbidity between severe cognitive impairment and major depression. There were only three cases of comorbidity between the two disorders; none of these cases died during the study period. Indeed, when we examined comorbidity between mild cognitive impairment and major depression, none of these additional seven individuals died during the study period. The reduction in effects of cognitive impairment and mortality need to be attributed to the older ages and the increased problems in physical health and functioning of this group.

The apparent lack of effect of other psychiatric disorders, especially substance abuse on mortality, needs to be viewed in the context of the age of the sample and the relatively short follow-up period. Substance abuse problems generally begin in youth, therefore the negative effects of physical health and mortality may be more apparent in younger ages. The surviving substance abusers in the age 55 and over group may be qualitatively different from others. A longer follow-up of adults of all ages which takes into account

age of onset of substance abuse problems would help clarify this issue.

In conclusion, the 15-month follow-up of 3,007 members of the New Haven community, ages 55 and over, indicates that risk of dying is increased for those with major depressive disorder, even controlling for other known risk factors. In future research on this sample, we hope to examine the longer term effects of psychiatric disorders on mortality and learn more about the process by which psychiatric and physical health factors interact to affect mortal risk.

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'Keep WIC Lit' Campaign to Help Hungry Women, Infants, and Children

Bread for the World, a Christian citizens' movement in the United States, has launched a campaign seeking full funding for the WIC program, one of the most effective government nutrition programs in the nation.

People are encouraged to mail a small birthday candle and a message—"Keep WIC Lit. Fully fund the WIC program so that more babies will celebrate their first birthday"—to their senators and representatives from now through August, to draw attention to the needs of millions of pregnant women, their infants and young children who are at nutritional risk.

The WIC program provides nutritious food, referrals to health services, counseling, and education to low-income women, new mothers, their infants and children up to age 5 years. The program saves US taxpayers millions of dollars a year; a Harvard study determined that every $1 spent on pregnant women in the WIC program saves $3 in hospital costs for newborns. The "Keep WIC Lit" campaign seeks increases of $150 million above inflation over each of the next five years in the WIC program so it can reach an additional 1.5 million nutritionally needy individuals. Less than half of those who are eligible are receiving WIC benefits at the present time.

For further information, contact: Katherine Smith, Bread for the World, 802 Rhode Island Avenue, NE, Washington, DC 20018. (Telephone 202/269-0200.)

Health Practices of the Elderly Poor

James E. Lubben, MPH, DSW, Philip G. Weiler, MD, MPH, and Iris Chi, DSW

Abstract: The purpose of the present study was to describe the health practices of the elderly poor and to examine the association between specific health practices and subsequent hospital use. Data came from a sample of 931 Medicaid elderly living in California. Smoking, limited social networks, and lack of regular exercise significantly increased the odds of subsequent hospital utilization. Implications of these findings, which could benefit health promotion intervention and treatment programs for the elderly, were discussed. (*Am J Public Health* 1989; 79:731–734.)

Introduction

There has been increasing attention paid to health promotion and disease prevention activities in the elderly because of economic, medical, and social concerns.[1,2] With increasing life expectancy, a focus on preventive measures to decrease morbidity and improve quality of life in old age has also developed.[3,4] To that end, health behavior and lifestyle have become important areas of concern over the last 20 years.

In the early 1970s, Belloc and Breslow[5] identified a list of seven health practices which were associated with health status: having never smoked; drinking less than five drinks at one sitting; exercising; sleeping seven to eight hours a night; maintaining desirable weight for height; avoiding snacks; and eating breakfast regularly. Inquiry regarding compliance with these practices has been included in the National Health Interview Surveys (NHIS) of the adult US population in 1977, 1983, and 1985.[6,7] Additional investigations have substantiated the importance of the smoking, drinking, exercise, and weight control health practices.[8–10] However, other practices have seldom been found to be significant.

More recently, other general health practices have also been suggested as important: maintaining a balanced diet (fruits, vegetables, and whole grains but limits on caffeine and salt intake) has been suggested as an important health practice,[1,11–13] and maintaining social networks.[9,14–20]

Current knowledge of specific health practices for elderly persons tends to be based on studies of younger populations. One of the few studies of an aged population reported that elderly who followed good health practices were not significantly advantaged in terms of mortality rates over a five-year period.[21] However, another study affirmed the value of some health practices for an elderly population when they examined mortality rates over a 17-year period,[10] finding that not smoking, getting regular exercise, maintaining weight control, and regularly eating breakfast were associated with lower mortality rates.

One aim of the present study was to evaluate the relative importance of specific health practices among elderly persons; another purpose was to explore differences in health practices of the elderly poor and the general elderly population.

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Methods

Design

The data were drawn from the California Senior Survey (CSS), a random sample of elderly Medi-Cal (Medicaid) recipients from eight representative communities within California. Eligibility for Medi-Cal includes those elderly on Supplemental Security Income (SSI) or with incomes of $550.00/month (single aged) or less after accounting for medical expenses. The CSS respondents were randomly selected from files which contain the names, addresses, and other basic demographic data for all Medi-Cal recipients. Only Medi-Cal recipients 65 years of age and older were sampled. CSS respondents were interviewed face-to-face in 1982 and paid $15.00 for their participation. Hospital utilization data were also obtained for one year following the interview. Comparisons between the survey data and statewide data indicated that respondents were typical of California's Medi-Cal elderly population. More complete descriptions of the sampling methodology and sample characteristics have already been published.[22–24]

For the present analyses, data from the interview were merged with those from other sources regarding a respondent's hospital use during the year following the interview. The hospital utilization data came from Medi-Cal paid claims computer tapes which were merged with data obtained from Medicare intermediaries.[25] Combining these two sources provided complete data on number of days hospitalized for 90 per cent of the original CSS sample. To assure accuracy, the computer tapes were examined for any bills submitted for hospitalizations during the follow-up year to either Medi-Cal or Medicare intermediaries on behalf of CSS respondents. To overcome any problems that might have been caused by late billing, bills submitted up to two years following the interview were reviewed. No significant differences in characteristics were noted between survey respondents with and without complete hospital data. Only respondents with complete hospital data were used for the following analyses (N=931).

The study relied upon self-report data regarding health practices using a questionnaire composed of several previously described instruments.[22–24] Many of the questions on health practices were taken verbatim from the original Belloc and Breslow study.[5]

Participants

Reflecting the diversity of California's ethnic population, the sample included a sizable proportion of Blacks (17 per cent), Hispanics (13 per cent), and Asians (12 per cent) although a majority (58 per cent) were White. With respect to gender, there were approximately two females for every male. The average age was 77.2 years and a large majority (61 per cent) of the respondents lived alone. With respect to health, 36 per cent rated their health as good, 41 per cent rated

it fair, and 22 per cent rated it poor. More than 75 per cent of the CSS respondents had seen a doctor within the past three months.

Health Practice Variables

Appendix A describes the inquiry used to construct the health practice variables analyzed in the present study. Belloc and Breslow[5] definitions of health practice for smoking, sleep, limited snacks, and regular breakfasts were used. Because alcohol consumption was so limited among the elderly in the present sample (68 per cent abstained), differences in this commonly examined health practice could not be evaluated.

Hospital Use Variables

Two dichotomous variables measuring hospital utilization were used: whether a respondent had been hospitalized within a year following the interview, and whether a respondent experienced a hospital stay of six or more days in the year following the interview. The figure of six days was selected because it was slightly higher than the average length of stay for all respondents (mean length of stay = 5.2 days).

Analysis

The present analysis involved a series of logistic regression models examining the relation of the 12 health practices with the two indices of subsequent hospital utilization. Controlling for age and self-reported health status at the baseline interview enabled us to compare our results with those from an earlier study which used a similar analytic model.[10]

Results

Prevalence of Risk Factors

Table 1 reports the percentage of the sample who did not comply with a specific health practice. Typically, 20 per cent to 25 per cent of the sample did not comply with any given health practice. Considerably higher non-compliance rates were observed for weight control (36 per cent), never smoking (43 per cent), regular physical exercise (50 per cent), and seven to eight hours of sleep (60 per cent). Only 12 per cent of the sample drank five or more cups a day of coffee or tea.

Subsequent Hospitalizations

Table 1 also shows how non-compliance with a specific health practice affected the odds of subsequent hospital use. The practices with the largest odds ratios were the two smoking items. Elderly who had limited social networks were half again as likely to be hospitalized and have extended stays as were elderly with larger networks. Similarly, a minimal level of physical activity reduced the odds of any hospitalization as well as the chances for an extended stay. Although almost all of the remaining odds ratios were in the expected direction, the 95 per cent confidence interval included 1.00. Separate analysis was conducted on underweight and overweight risk factors but neither was related to hospital use. Thus, based upon the present study, the most important health practices for elderly persons appeared to be not smoking (ever or currently), maintaining social networks, and engaging in some form of physical activity at least once a week.

Since a major illness could have both changed an elderly person's health practices and also caused a hospitalization, the logistic regression models were modified to include a dummy variable indicating whether a respondent had been hospitalized in the year prior to the interview. The results of these additional analyses showed that maintaining social networks (OR 1.52; CI=1.08 2.12), not ever smoking (OR 1.59; CI=1.15, 2.21), and not currently smoking (OR 1.59; CI=1.08, 2.33) continued to be related to risk of any subsequent hospitalization. However, the importance of regular physical exercise was diminished (OR 1.29; CI=.95, 1.76). Similar results were observed when considering the risk of a stay of six days or more in a hospital: maintaining social

TABLE 1—Association between Behavioral Risk Factors and One Year All Cause Hospitalization Rates

Behavioral Risk Factors	Per Cent at Risk	Hospitalized		Hospitalized for at Least 6 Days	
		Odds Ratios[a]	95% CI	Odds Ratios[a]	95% CI
Smoking					
Ever	43	1.59	1.16, 2.18	1.76	1.23, 2.50
Current	21	1.56	1.07, 2.27	1.68	1.11, 2.52
Social Networks	23	1.51	1.09, 2.10	1.57	1.10, 2.24
Limited networks					
Physical Activities	50	1.43	1.05, 1.94	1.52	1.08, 2.14
Less than weekly					
Whole Grains	25	1.15	.82, 1.63	1.41	.97, 2.05
<2 servings/day					
Weight	36	1.07	.79, 1.46	1.31	.93, 1.83
10% under or 30% overweight					
Fruits & Vegetables	22	1.18	.83, 1.68	1.26	.86, 1.86
<2 servings/day					
Sleeping Patterns	60	1.26	.93, 1.69	1.16	.83, 1.62
<7 or >8 hours/day					
Caffeine	12	1.21	.77, 1.92	1.12	.67, 1.88
5 or more cups/day					
Salt	20	1.00	.68, 1.45	1.04	.69, 1.58
Uses table salt or eats salty foods daily					
Snacks	22	.97	.69, 1.37	1.00	.68, 1.47
Sometimes or often					
Breakfast	18	.74	.50, 1.10	.92	.60, 1.40
Does not eat breakfast regularly					

[a]Odds ratios are adjusted for age and self-reported health at baseline.

networks (OR 1.57; CI=1.10, 2.26), not ever smoking (OR 1.76; CI=1.23, 2.53), and currently smoking (OR 1.70; CI=1.12, 2.58), were significant risk factors whereas regular physical exercise (OR 1.38; CI=.97, 1.95) was not.

Discussion

About 10 per cent of California's elderly are Medicaid beneficiaries, compared with a national average of 8.3 per cent. Thus California's elderly Medi-Cal recipients may be fairly similar to elderly Medicaid recipients in general. However, the Medicaid elderly respondents in the present study are only indicative of the poorest of the elderly poor. Many elderly with limited incomes are unable to qualify for Medicaid because of other assets such as a savings account in excess of $2000. Thus, only one-third of all US elderly persons below the poverty line receive Medicaid benefits.[26]

The present findings are subject to the limitations of all cross-sectional studies of health practices and so the reported relationships between health practices and hospitalization should be viewed as associational and not causal. However, the magnitude of the increased odds of hospitalization associated with smoking were similar to previously reported levels of association between smoking and mortality rates.[10,27] Further, the relation between social networks and hospital use in the present study was also similar to that reported between social networks and mortality rates.[9]

The present findings failed to support some health practices found significant in earlier studies. For example, weight control and regular breakfast were previously reported to be significantly related to mortality rates[10] but neither were related to hospital rates in the present study. Different dependent variables (hospital vs mortality) may have accounted for this difference in results. None of the dietary health practices were highly associated with hospital use but our measures were not very sensitive.

The prevalence of smoking, lack of exercise, and weight problems was particularly high among the California elderly Medicaid recipients. Since three-fourths of our sample had seen a doctor within the last six months, it would seem most appropriate that these health practices receive increased emphasis by physicians serving the elderly poor. However, increased physician attention to preventive medicine may require a re-examination of public and private health insurance coverage for the elderly. Somers argues this point very well by stating it makes no sense to reimburse a physician for treating an illness but not for preventive care that maintains health.[28] Because public funds are the primary source of support of health care for the elderly poor, public policy must address the issue of financing health promotion for this population.

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APPENDIX A
Health Practices Variables used in the Analysis

Health Practice:	Questionnaire Item	Favorable Response
Smoking		
Ever	Have you ever smoked cigarettes regularly?	No
Current	Do you smoke cigarettes at the present time?	No
Social Networks	Series of questions covering following aspects of social networks: 1. Frequency of contact with a relative 2. Number of relatives seen monthly 3. Number of relatives feels close to 4. Frequency of contact with a friend 5. Number of friends seen monthly 6. Number of friends feels close to 7. Has someone to talk to about important decisions 8. Others talk to respondent when they have important decisions 9. Feels relied upon and helps others 10. Lives with others	Above lower quartile on the Lubben Social Network Scale[a]
Physical Activity	How often do you exercise, for example do calisthenics, vigorous gardening, or take long walks?	Weekly or more often[b]
Whole Grains	How many servings of bread, cereal, or whole grains do you eat in a day?	Two or more
Weight Control	What is your weight? How tall are you?	Between 10% underweight and 30% overweight[c]
Fruits and Vegetables	How many servings of fruits and vegetables do you eat in a day?	Two or more
Sleep Patterns	How many hours of sleep do you usually get at night?	Between 7 and 8 hours
Caffeine	On an average day, how many cups of coffee and tea do you drink?	Less than five[d]
Salt	How often do you eat salty type foods such as potato chips, pickles, lunch meats or canned soups?	Once in a while, rarely or never
	How often do you add salt to your food at the table?	Once in a while, rarely or never
Snacks	How often do you eat in between your regular meals?	Rarely or never
Breakfast	Do you eat breakfast most mornings?	Almost every day

[a]Lubben[29] modified the original Berkman-Syme[14] Social Network Index specifically for use with older populations. Cronbach alpha for the 10 items included in the LSNS was .70.
[b]Due to age of respondents, lower levels of exercise were tolerated than in original Belloc-Breslow study.[5]
[c]The 1960 Metropolitan Life Tables[30] were used to conform with recent NCHS data.[6] The standards were also similar to those used in other studies of elderly persons.[10,31]
[d]Standard suggested by LaCroix, *et al.*[32]

APHA Issues Report on "HIV Antibody Testing"

HIV Antibody Testing—the 4th report of the Special Initiative on AIDS, a program activity of the American Public Health Association—has been released. This new report, prepared by the AIDS Working Group, summarizes the use and the purpose of the human immunodeficiency virus (HIV) antibody tests. The necessary components and considerations of testing programs are reviewed and proposals to screen specific populations are discussed.

To date four reports have been issued in this AIDS Report Series.

- No. 1, *Casual Contact and the Risk of HIV Infections* was released in July 1988
- No. 2, *Contact Tracing and Partner Notification*, November 1988
- No. 3, *Illicit Drug Use and HIV Infection*, January 1989
- No. 4, *HIV Antibody Testing*, April 1989.

Single copies of the various reports are available for $3 each prepaid from APHA Publication Sales, 1015 15th Street, NW, Washington, DC 20005. Bulk rates available on request.

Mechanical Restraints on the Medical Wards: Are Protective Devices Safe?

Richard P. Lofgren, MD, MPH, David S. MacPherson, MD, Rosanne Granieri, MD, Sharon Myllenbeck, RN, and J. Michael Sprafka, PhD, MPH

Abstract: We prospectively identified 102 mechanically restrained hospital patients and determined their hospital course. The cohort was elderly, cognitively impaired with multiple chronic diseases. The in-hospital mortality was 21 percent. Nosocomial infection developed in 12 percent and new pressure sores in 22 percent. Prolonged use of restraints (>4 days) was the strongest independent predictor of nosocomial infection (relative risk 1.8, 95% CI = 1.2, 2.8) and new pressure sores (RR 1.4, 95% CI = 1.1, 1.8) as determined by multiple logistic regression analysis. Patients placed in mechanical restraints for longer than four days experience frequent morbid events and should be monitored carefully. (*Am J Public Health* 1989; 79:735–738.)

Introduction

Mechanically restrained patients are a familiar sight on the medical wards in many hospitals in the United States. A survey of a general hospital found that more than 7 percent of patients admitted to the general medical floors were restrained at some time during their hospital stay.[1] Elderly veterans were found in restraints in 17 percent of admissions to medical wards.[2] By contrast, restraints are used much less frequently in other countries such as the United Kingdom.[3,4]

Posey vests, ankle and wrist restraints are often used to bridle confused patients for the purpose of preventing falls. Between 70 to 80 percent of all inpatient incident reports are related to patient falls.[5–9] Accidents and falls do occur despite the use of these devices.[10] A recent study comparing two Canadian hospitals found that the institution that used restraints less often had about half as many accidents.[3] Little is known about possible complications related to the use of mechanical restraints. Patients have been inadvertently hung and strangled to death trying to escape from restraining devices[11,12] and episodes of hyperpyrexia and rhabdomyolsis have been reported.[13,14]

We prospectively followed a cohort of mechanically restrained patients residing on the general medical service of an acute care hospital to describe the characteristics and clinical outcomes of this population of patients. We further identified risk factors for death, new pressure sores and nosocomial infections.

Methods

We prospectively followed all patients residing on the general medical wards, who were initially placed in mechanical restraints during a 13-week enrollment period. Patients in the intensive care, coronary care, special diagnostic study, and restorative care units were excluded. Mechanically restrained patients were identified by the charge nurse each shift. The protocol was approved by the hospital's human studies committee and informed consent was obtained from each patient or next of kin.

Within 48 hours after patients were first restrained, one of the investigators reviewed the medical record and examined the patient. Data were collected using a standard form. Demographic and clinical data were abstracted from the chart using the housestaff and nursing notes. Administration of medications was determined through nursing records. Resuscitation status was determined through review of orders and housestaff notes. Each patient was examined initially for the presence of pressure sore, incontinence, and indwelling medical devices. A mini-mental status examination was performed,[15] the level of consciousness was estimated, and an assessment of the patient's ambulatory and functional status was determined.[16]

Throughout the hospital stay, the record of each mechanically restrained patient was reviewed every 48 hours and the patient re-examined with attention to nosocomial infections, status of pressure sore, presence of incontinence, and presence of indwelling medical devices. Falls were considered to have occurred if noted by nursing or physician staff or if noted in an incident report form. At discharge, the length of stay and disposition of the patient was noted.

A mechanical restraint was defined as any device applied to a patient for the purpose of limiting free movement, such as a posey vest, wrist restraint, or full sheet restraint. This did not include bed side rails or geri-chairs. Pressure sore was defined as a local area of epidermal defects or necrosis over pressure points or bony prominence ranging from grade I (reactive hyperemia lasting more than 24 hours) to grade V (deep ulceration extending to the bone).[17] A nosocomial infection was defined as one that was neither present nor incubating at the time of the patient's admission and had its onset during hospitalization. New fecal or urinary incontinence was noted if soilage was found on physical examination, a condom or indwelling urinary catheter was present, or a notation by a physican or nurse was found in the record. Infusion thrombophlebitis was considered to be present if two or more of the following signs and symptoms were present: pain, erythema, swelling, induration, or palpable venous cord.

Descriptive statistics were calculated for the cohort of restrained patients. Demographic and clinical characteristics were examined to determine which factors were associated with adverse outcomes, including death, development of pressure sore, nosocomial infection. Stepwise logistic regression analysis was used to determine factors independently and significantly associated with adverse outcomes. Statistical analyses were performed using SPSS[19] and BMDP[20] statistical software.

Results

Of the 1,661 patients admitted to the general medical wards during the study period, 102 (6 percent) were mechan-

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ically restrained at some point during their hospital stay. The restrained patients were elderly with an average age of 75.4 years (s.d. = 12.7, range 50 to 98 years). All but one were male. Restrained patients were admitted from home (42 percent), nursing homes (27 percent), transfers from other hospitals (15 percent) or intensive care units (13 percent), or elsewhere (3 percent). Over half of the patients were restrained on the day of admission. The average time from admission to when the patient was first restrained was 1.8 days (range from 0 to 17 days). Patients were entered into the study at the time they were first restrained and observed for the remainder of their hospital stay. The average length of stay in the hospital was 13.4 days (±13.3 days). Posey vests were used in 61 percent of patients, soft wrist restraints in 14 percent, and both in 24 percent.

The clinical characteristics of the restrained cohort are summarized in Table 1. The most common admitting diagnoses fell into three major categories: infectious diseases (25 percent); central nervous system (CNS) disorders (structural or metabolic) (20 percent); and cancer (15 percent). Almost all the patients were cognitively impaired.

Adverse events were common among the restrained patients (Table 2). Nearly 60 percent of the patients admitted from home were discharged to chronic care facilities. Eighty-two different pressure sores were seen in 37 of the 102 mechanically restrained patients; 36 were new sores that developed in 22 of the patients. Four patients developed sores that progressed to grade II or greater. Of the 82 sores visualized, 63 percent were grade I, 29 percent were grade II, and 8 percent grade III. During the course of hospitalization, 23 percent of the pressure sores increased in size, 26 percent decreased or resolved, and 51 percent did not change.

Different demographic and clinical characteristics were analyzed for associations with morbid events (Table 3). Nosocomial infections and new pressure sores were most strongly associated with prolonged use of mechanical restraints (dichotomized by the upper tercile of length of time in restraints; >4 days). In-hospital mortality was associated with a DNR (do not resuscitate) status and nosocomial infection.

Using a logistic regression model, we investigated the extent to which various clinical parameters might affect the probability of an adverse outcome. For each dependent variable, nine independent variables were entered in the model (Table 4). Consistent with the univariate analysis, the length of time restrained was a significant predictor of nosocomial infection and new pressure sores but not of in-hospital mortality.

TABLE 2—Summary of the Clinical Course of the 102 Restrained Patients during Their Hospitalization

Outcome	Percent
Falls	4
New Urinary Incontinence*	29
New Fecal Incontinence*	29
Infusion Thrombophlebitis**	13
New Pressure sore	22
Nosocomial Infection	12
In-Hospital Mortality	21

*This represents percentage of patients who were initially continent who became incontinent after being placed in mechanical restraints.
**105 different intravenous devices were used in this cohort. Twelve were in place >72 hours and there were no cases of septic phlebitis.

To adjust for possible differences in total exposure time (or length of stay) between the patients restrained for more than 4 days (long-time group, N = 35) and those restrained for 4 days or less (short-time group, N = 67), we examined the rate of the morbid events per patient-hospital day. The rate of nosocomial infections and new pressure sores were higher in the long-time versus short-time restrained group: 26 versus 3 nosocomial infections per 1,000 patient-hospital days (95% CI of the difference = 3, 42) and 38 versus 13 new pressure sores per 1,000 patient-hospital days (95% CI of the difference = 1, 49). Consistent with the multivariate analysis, other indirect measures of severity of illness were similar between the two groups including number of chronic diseases, discharge diagnoses, mobility status, DNR status, use of intravenous lines, and foley catheters.

Discussion

Over 6 percent of patients residing on the general medical floors at our institution were restrained at some point during their hospitalization. Mechanically restrained patients were usually elderly and often confused with multiple chronic diseases. Nearly one-half of these patients received some

TABLE 1—Clinical Characteristics of 102 Mechanically Restrained Patients

Underlying Conditions	Mean ± SD	Median	Percent
Number of Chronic Diseases	4.6 (±0.2)		
Number of Discharge Diagnoses	6.8 (±0.3)		
Number of Medications Prescribed at Time of Restraints	4.6 (±0.2)		
Do Not Resuscitate (DNR) Status			57
Use of Centrally Acting Medications*			49
Physical Examination			
Mini-Mental Status Score (median)**		10.8	
Ambulates Independently			13
Pressure sore Present at Time of Restraints			28
Incontinence Prior to Being Restrained			
Urine			46
Fecal			23
Devices Used during Hospitalization			
Foley Catheter			29
Intravenous Lines			63
Nasogastric Tube			12

*The centrally acting medications used in this cohort included narcotics (25%), benzodiazepam (35%), and phenothiazines (19%). Some patients receive more than one of these medications thus the total is >49%. **A mini-mental score of <24 suggests "probable" dementia and a score <15 indicates "definite" dementia. Only 9.8% of the patients had a score >24.

TABLE 3—Clinical Factors Associated with Untoward Outcomes*

Outcome	Parameters	Odds Ratio	95% CI
In-Hospital Mortality	DNR Status	23.3	3.0, 182
	Nosocomial Infection	4.3	1.1, 16.0
New Pressure sore	Length of Time Restained	6.3	2.0, 20.3
	Previous Pressure sore	3.8	1.2, 11.5
	Nosocomial Infection	4.4	1.1, 18.4
	New Fecal Incontinence	2.2	0.7, 7.0
Nosocomial Infection	Length of Time Restained	12.4	2.5, 60.7
	Length of Hospital Stay	7.5	1.6, 36.4
	Nasogastric Tube	4.9	1.2, 20.1
	Intravenous Lines	6.8	0.8, 55.5

*Clinical factors included length of time restrained, length of stay, age, new or previous pressure sore, new or previous fecal incontinence, new or previous urinary incontinence, mini-mental score, ambulation status, DNR status, number of chronic illness, number of discharge diagnoses, number of medications, number of centrally acting medications, condom or foley catheters.

TABLE 4—Significant Predictors of Adverse Outcomes among Mechanically Restrained Patients Using a Logistic Regression Model

Outcome	Predictors	Relative Risk	95% C.I.
In-Hospital Mortality*	DNR Status	53.7	3.0, 956
	Nosocomial infection	26.7	2.6, 268
	No. Chronic Diseases	1.4	0.9, 2.1
Nosocomial Infection**	Length Restrained	1.8	1.2, 2.8
	Mini-Mental Score	1.2	0.9, 1.3
New Pressure Sore†	Length Restrained	1.4	1.1, 1.8
	Age	1.05	0.99, 1.1

*The dependent variables included in this model were length of time restrained, age, mini-mental score, ambulation status, DNR status, number of chronic illness, number of medications, nosocomial infection, use of foley catheters

**The dependent variables included in this model were length of time restrained, age, previous pressure sore, fecal incontinence, mini-mental score, number of chronic illness, number of medications, and use of foley catheters or intravenous lines.

†The dependent variables included in this model were length of time restrained, age, previous pressure sore, new or previous fecal incontinence, new urinary incontinence, mini-mental score, ambulation status, and number of chronic illnesses.

medication that would affect mental capacity. Undesirable clinical outcomes were frequent in this group of patients. The mortality rate of restrained patients was quite high with one of five not surviving hospitalization. Sixty percent of those admitted from home were discharged to a chronic care facility.

The use of mechanical restraints at our institution is similar to other reports. Frengley, *et al*, surveyed all patients admitted to the medical service of a large urban city/county hospital in Cleveland and found that 7.4 percent were restrained.[1] Robbins, *et al*, studied all patients greater than 70 years of age and a random sample of those less than 70 years of age admitted to the Denver VA Hospital and found and that 17 percent were restrained during hospitalization.[2] The incidence of use in our study was lower, in part, because we did not include patients restrained in geri-chairs. Both of the other studies noted an 8- to 12-fold increase in mortality among the restrained population. These studies, however, did not examine other clinical outcomes.

Our cohort of restrained patient suffered extraordinarily high rates of in-hospital death and morbid events (Table 2). Patients restrained for greater than 4 days developed significantly more nosocomial infections and new pressure sores when compared to the group restrained for the shorter period. Using a multiple logistic regression model, the length of time restrained was found to be the only independent predictor of nosocomial infection and new pressure sore. Mortality was unaffected by the duration of restraint.

The increased observed morbidity noted in the long-time restrained group may be due to a difference in severity of illness. However, the clinical characteristics of the two groups were quite similar (Table 3). With our logistic regression model, we attempted to adjust for measures of potential difference in severity of illness (such as the number of discharge diagnoses). Although the groups appeared to be clinically comparable, it is possible that subtle differences still existed.

There are several other potential limitations to this study. The patient population was a homogeneous group of elderly men hospitalized on the general medical wards of a large urban VA medical center and conclusions regarding restraint use must be applied with caution to other clinical situations such as community hospitals, psychiatric units, intensive care units, or extended care facilities. Our cohort study only consisted of patients who were restrained and did not contain a clinically comparable group that was never restrained. We are therefore unable to establish a causal relationship between the use of mechanical restraints and morbid events. To firmly establish a causal relationship, a controlled trial in which eligible patients were randomly assigned to receive mechanical restraints versus some alternative therapy would be necessary. In the meantime, our findings demonstrate an important dose effect; those patients restrained for longer periods of time had more adverse outcomes.

It is not clear why prolonged use of restraints is associated with morbid outcomes. Restraining patients may be a form of medical and social isolation. It is more difficult to turn or examine patients while they are in restraints. Essential patient care procedures may therefore not be routinely

performed. Restraints may cause nurses and physicians to feel patients are unable to communicate. Loss of autonomy and social isolation have been shown to adversely affect the physical and mental health of elderly patients.[21,22] Less intense care and isolation may increase the chance of an adverse event.

The results of our study suggests that mechanically restrained patients need to be closely monitored. If mechanical restraints seem necessary, effort should be made to limit their use to no more than four days. Further investigations are needed to confirm our findings and to identify the reasons that morbid events are increased with the use of restraints. Alternate safety measures, free of associated morbidity, should be developed to protect this vulnerable population.

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Cost-Effectiveness of Antibiotic Prophylaxis for Dental Procedures in Patients with Artificial Joints

JOEL TSEVAT, MD, MPH, ISABELLE DURAND-ZALESKI, MD, MPH, AND STEPHEN G. PAUKER, MD

Abstract: We performed a cost-effectiveness analysis to evaluate whether patients with artificial joints should take penicillin, erythromycin, or no antibiotics before dental procedures. We modeled the risk of anaphylaxis from penicillin, the risks and consequences of an artificial joint infection, and the actual variable costs of hospitalization and antibiotics.

Penicillin prophylaxis is slightly less expensive than erythromycin prophylaxis but is both more expensive and less effective than no prophylaxis. Erythromycin prophylaxis, the most effective, is the most expensive strategy. The marginal cost effectiveness of erythromycin prophylaxis compared to no prophylaxis is $12,900 per quality-adjusted year of life saved. Sensitivity analysis demonstrates that the risk of developing a joint infection is the key parameter in the analysis.

Based on our estimated risk of developing a joint infection, the cost-effectiveness of antibiotic prophylaxis with erythromycin compares favorably with other medical interventions. Thus, until a definitive study to quantify the risk is conducted, patients with artificial joints should take prophylactic erythromycin when they undergo dental procedures. (*Am J Public Health* 1989; 79:739–743.)

Introduction

Antibiotic prophylaxis to prevent bacterial endocarditis is accepted medical practice in patients with valvular heart disease who undergo dental manipulation.[1] Whether or not patients with prosthetic joints should take prophylactic antibiotics to prevent septic arthritis has not been established. Dental infections have been shown to cause septic arthritis in artificial hips via hematogenous spread,[2] but few if any cases of prosthetic joint infections directly attributable to dentally induced transient bacteremias have been reported and confirmed.[3]

Nevertheless, many authorities recommend antibiotic coverage during dental procedures for patients with artificial joints.[4–6] Jaspers and Little found that although only 41 per cent of orthopedic surgeons believed that dentally induced bacteremias could cause prosthetic joint infections, 93 per cent recommended giving antibiotics to their patients.[3]

The scope of the problem is large: in the United States in 1986, there were 201,000 total hip replacements and other hip arthroplasties, 91,000 total knee replacements, and 245,000 other upper and lower extremity arthroplasties and joint replacements—a total of 537,000 arthroplasties.[7] The risks of administering prophylactic antibiotics primarily involve adverse reactions including anaphylaxis. The potential benefits, on the other hand, are infections and hospitalizations averted and lives saved. The issues then are both clinical and economic. We explored these issues in a cost-effectiveness analysis.

Methods

The Decision Model

The problem can be represented as a decision tree (Figure 1). The model encompasses only one visit to the dentist and its consequences. The three branches emanating from the square decision node represent the three strategies we examined: the dentist can prescribe either penicillin or erythromycin or choose to give no prophylaxis. Ensuing events are beyond the dentist's control and are represented as round chance nodes.

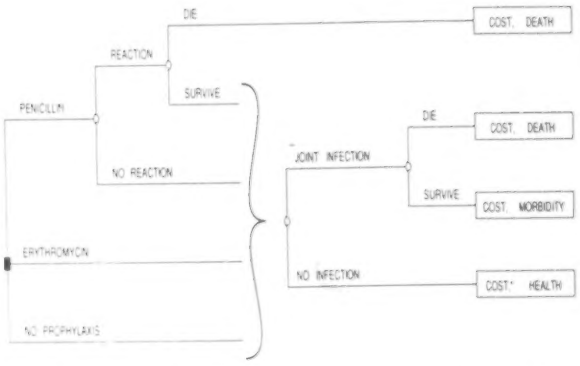

FIGURE 1—The Decision Tree (cost = 0 for the no infection branch of the no prophylaxis strategy)

Assumptions

- Infections resulting from transient bacteremia during dental procedures would be caused exclusively by oral flora, namely by streptococci (including anaerobic streptococci). Therefore, we consider only the 19 per cent of late infections (more than six months postoperative) of hip and knee prostheses caused by *Streptococcus spp*.[4]
- Because there are no data concerning the effectiveness of antibiotics in preventing artificial joint infections resulting from dental procedures, we assume that the effectiveness of prophylactic therapy is the same as for the prevention of bacterial endocarditis in patients with valvular heart disease.
- Because the elderly comprise a large proportion of patients with artificial joints, we derive the probability of developing a joint infection during one visit to the dentist by using the average number of dental visits per year for a cohort of 65-year-olds.
- Only dentulous patients are at risk for transient bacteremias from dental procedures. Consequently,

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we exclude the 40 per cent of elderly patients who wear dentures from the analysis (Athena Papas, DDS, personal communication).
- Minor medication side-effects (rash and gastrointestinal symptoms) incur no monetary treatment costs but do incur short-term morbidity.

Probabilities

Probability of developing an artificial joint infection from a transient bacteremia during a dental manipulation: The risk for dentulous patients is 9.3×10^{-5} or 9.3/100,000 dental procedures (see Appendix).

Mortality from an infected artificial joint: Little found that 18 per cent of patients with acute dental infections who had hematogenously infected hip prostheses died.[2]

Anaphylactic reactions: The probability of having an anaphylactic reaction to penicillin has been variably estimated to be between 0.0001 (1/10,000)[8] and 0.0004 (1/2,500).[9] In this analysis we use 0.0002 (1/5,000). If one in 50,000 (0.002 per cent) patients taking penicillin has a fatal anaphylactic reaction,[10] then one in 10 patients having an anaphylactic reaction dies.

We assume that erythromycin does not cause anaphylaxis.

Effectiveness of prophylactic antibiotics: The effectiveness of prophylactic penicillin in preventing septic arthritis is defined as:

$$1 - \left(\frac{\text{the probability of developing septic arthritis in a treated patient}}{\text{the probability of developing septic arthritis in an untreated patient}} \right)$$

A recent case-control study[11] found the effectiveness of prophylaxis in patients with valvular heart disease to be 0.91; the study did not differentiate among antibiotic regimens, however. The effectiveness of prophylactic penicillin in preventing bacterial endocarditis in patients undergoing dental procedures has been estimated by a panel of experts[12] to be 90 per cent. The panel estimated the effectiveness of erythromycin to be 75 per cent. We employ those estimates in this analysis.

Utilities

Possible outcomes are **health** (no infection), **death** (from anaphylaxis or joint infection), and **morbidity** from an infection. Outcomes are expressed as "quality-adjusted life expectancy" for a prototypic 65-year-old patient. The unadjusted life expectancy is derived from 1981 US life tables.

Death: Death is assigned a utility of zero.

Health: A year of health is assigned a utility of one full quality-adjusted life year (QALY).

Morbidity from a joint infection: A year during which a patient is hospitalized for an infected artificial joint is assigned a utility of 0.9 quality-adjusted life years.

Short-term morbidity from penicillin and erythromycin prophylaxis: For patients who do not die from a joint infection or from anaphylaxis, we deducted approximately 8 minutes from their quality-adjusted life expectancy for a course of prophylactic antibiotics (see Appendix).

Costs

Because this analysis takes the perspective of a health maintenance organization (HMO), actual costs rather than charges are used.[13] All costs were converted to 1986 dollars using a 5 per cent annual inflation rate. Future costs are not discounted against present costs because "future costs" (costs of treating joint infections) are incurred only a few months after the visit to the dentist.

Cost of antibiotics: The actual acquisition costs to one large metropolitan hospital's pharmacy of six 500-mg penicillin V potassium tablets and six 250-mg tablets of erythromycin are 28 cents and 30 cents, respectively (E Anderson, Jr, MS, personal communication). We used the recommended regimen for patients with cardiac conditions: penicillin V 2.0 g orally one hour before the procedure and 1.0 g orally 6 hours later, or erythromycin 1.0 g orally one hour before followed by 500 mg 6 hours later.[1]

To estimate the total variable cost of the antibiotics, we added a dispensing cost of $3.43, the average of a price bid by a mid-sized pharmacy chain for a contract with an HMO ($3.09 per prescription) and the internal cost to a large HMO with a house pharmacy ($3.76 per prescription) in 1986 (NM Kane, DBA, personal communication).

Cost of artificial joint infections: Based on experience with all patients (six) at our hospital with late infections (beyond the immediate post-operative period) of hip or knee prostheses (October 1, 1984 to September 30, 1986), the mean marginal (variable) cost of hospitalization was $16,000. (The average total cost of hospitalization was $23,000.)

Cost of anaphylaxis: The mean marginal cost of all patients treated for anaphylaxis during the same time period was $1,200 (the mean total cost was $2,900).

Results

No prophylaxis is the least costly strategy; penicillin is the next, being just slightly less expensive than erythromycin. Erythromycin is the most effective strategy followed by no prophylaxis and then penicillin (Table 1). Thus penicillin is both more expensive and less effective than no prophylaxis. The marginal cost-effectiveness of erythromycin compared to no prophylaxis is $12,900 per quality-adjusted year of life saved.

Sensitivity Analyses

We varied each of the probabilities over the full range of possible values and the costs over a wide range (Table 2).

TABLE 1—Baseline Results

Strategy	Cost ($)	Effectiveness (QALY)	Marginal Cost ($)	Marginal Effectiveness (QALY)	Marginal Cost/Effectiveness ($/QALY)
No Prophylaxis	1.4938	16.6997	—	—	—
Penicillin	4.0944	16.6996	2.6006	−0.0001	Dominated
Erythromycin	4.0965	16.6999	2.6046	0.0002	12,879.69

QALY = quality-adjusted life years.

TABLE 2—Sensitivity Analyses

Variable	Value	Least Costly Strategy (LCS)	Most Effective Strategy (MES)	Marginal Cost-Effectiveness of MES Compared to LCS ($ Per Quality-Adjusted Life Year)*
Probabilities				
Joint infection	0	NO RX	NO RX	NO RX DOM
	0.00001–0.00002	NO RX	ERY	393,600–107,900
	0.00003–0.00004	NO RX	ERY	60,700–41,300
	0.00005–0.00008	NO RX	ERY	30,700–16,100
	0.000093†	NO RX	ERY	12,900
	0.00010–0.00025	NO RX	ERY	11,600–1,300
	0.00030–0.00050	PCN	ERY	2,600–9,600
	0.00055–0.00065	PCN	ERY	13,900–40,700
	0.00070	PCN	ERY	149,700
	0.00073+	PCN	PCN	PCN DOM
Case-fatality rate of a joint infection	0	NO RX	NO RX	NO RX DOM
	0.05–0.15	NO RX	ERY	51,000–15,600
	0.18†–1	NO RX	ERY	12,900–2,300
Anaphylactic penicillin reaction	0	NO RX	PCN	9,600
	0.00002	NO RX	PCN	11,200
	0.00005–0.0001	NO RX	ERY	12,900
	0.0002†–0.0005	NO RX	ERY	12,900
Anaphylaxis case-fatality rate	0	NO RX	PCN	10,600
	0.05	NO RX	ERY	12,900
	0.1†–1	NO RX	ERY	12,900
Effectiveness of penicillin prophylaxis	0–0.8	NO RX	ERY	12,900
	0.9†–1	NO RX	ERY	12,900
Effectiveness of erythromycin prophylaxis	0	NO RX	NO RX	NO RX DOM
	0.1–0.7	NO RX	ERY	241,100–14,300
	0.75†	NO RX	ERY	12,900
	0.8–1	NO RX	ERY	11,700–8,100
Utilities				
Quality-adjustment factor for year with a joint infection‡	0–0.8	NO RX	ERY	10,300–12,500
	0.9†–1.0	NO RX	ERY	12,900–13,300
Short-term morbidity from antibiotic prophylaxis	0	NO RX	ERY	12,000
	7.5 mins.†	NO RX	ERY	12,900
	60 mins.	NO RX	ERY	25,500
Costs (1986 $)				
Artificial joint infection	0–15,000	NO RX	ERY	18,400–13,200
	16,000†–30,000	NO RX	ERY	12,900–8,000
Anaphylaxis	**0–1,200†**	NO RX	ERY	12,900
	2,000+	NO RX	ERY	12,900
Penicillin acquisition cost¶	**0–0.28†**	NO RX	ERY	12,900
	0.50+	NO RX	ERY	12,900
Erythromycin acquisition cost¶	**0–0.30†**	NO RX	ERY	11,400–12,900
	0.50–2.50	NO RX	ERY	13,900–23,800
Antibiotic dispensing cost	0–0.06 (20%)	PCN	ERY	14
	1.00–2.00	NO RX	ERY	900–5,900
	3.00–**3.43†**	NO RX	ERY	10,800–12,900
	4.00–5.00	NO RX	ERY	15,700–20,700

NO RX indicates the no prophylactic antibiotics strategy
ERY indicates the erythromycin prophylaxis strategy
PCN indicates the penicillin prophylaxis strategy
DOM indicates that the particular strategy is dominant (least expensive and most effective)
*Ratios greater than 100 are rounded to the nearest 100
†Value used in baseline analysis
‡Possible values range from 0 (equivalent to death) to 1 (equivalent to health prior to the infection), with intermediate values corresponding to "fractions" of "full health"
¶Excludes dispensing cost

Although the ranking of strategies by cost shifts with minor deviations from the baseline values of many variables, erythromycin tends to remain the most effective. For example, if the actual probability of an artificial joint infection is greater than 27/100,000, penicillin becomes the least expensive strategy; however, it remains less effective than erythromycin unless the risk of infection exceeds 72/100,000. The sensitivity analysis on the probability of joint infection subsumes a sensitivity analysis on the fraction of hematogenous infections caused by transient bacteremias from dental manipulations, assumed in the base case to be 10 per cent. Thus, all other things being equal, if the fraction of hematogenous infections due to dental manipulations exceeds 29 per cent, erythromycin becomes the least expensive strategy, but it remains less effective than erythromycin unless the fraction exceeds 77 per cent.

An important variable is the probability of anaphylaxis from penicillin, because the risk from oral administration may be lower than the published risk[8–10] in which the route of administration was not specified. Even if the incidence of allergic reaction is half as great after oral administration as after parenteral administration,[14] penicillin is less effective than erythromycin and the marginal cost-effectiveness of erythromycin is $12,900 per additional life saved compared to no prophylaxis. Penicillin becomes the most effective strategy (but remains more expensive than no prophylaxis) if the risk of fatal anaphylaxis is less than approximately 1/400,000 (baseline risk: 1/50,000).

The analysis is also very sensitive to outpatient antibiotic *dispensing* costs. If dispensing costs amount to only 20 per cent of acquisition costs (MC Weinstein, PhD, personal communication), or six cents per prescription, then penicillin is the least expensive (and least effective) strategy and the marginal cost-effectiveness of erythromycin compared to penicillin is only $14 per QALY saved. With dispensing costs of $3.76 per prescription, on the other hand, no prophylaxis is the least expensive strategy and the incremental cost of prescribing erythromycin is $14,500 per additional QALY gained.

Discussion

Based on our assumptions, erythromycin prophylaxis for dental procedures in patients with artificial joints seems to be worthwhile. Prophylaxis with penicillin, however, is both more expensive and less effectiveness than no prophylaxis. Several issues merit further consideration.

This cost-effectiveness analysis was performed from the perspective of an HMO. Only hospitalization and drug costs were considered. Because infected patients would have more frequent follow-up visits than the non-infected ones, prophylaxis with erythromycin could appear to be even more cost-effective if one includes the cost of outpatient follow-up. From the patient's perspective, non-medical direct costs (e.g., transportation) and indirect costs (e.g., lost wages) require consideration; including these would also bolster the case for erythromycin.

The cost of *dispensing* antibiotics in the outpatient pharmacy has a major impact on the analysis. In the base case, we assigned a dispensing cost ($3.43 per prescription) that exceeds the acquisition cost by over 1000 per cent! Perhaps the average dispensing cost exaggerates the cost of dispensing only six tablets, and a fixed fraction (e.g., 20 or 40 per cent) of the acquisition cost might be more appropriate. There is an opportunity cost to filling inexpensive prescriptions, however, and time spent filling an inexpensive prescription could have been spent filling a more expensive prescription. From an opportunity cost standpoint then, perhaps the real cost of dispensing the antibiotic is about $3.43. If so, the no prophylaxis strategy is the least expensive because dispensing costs dominate the costs of treating joint infections.

The costs of each strategy can be broken into their components (Table 3). In the no prophylaxis strategy, drug and anaphylaxis costs are zero; the entire cost consists of infection costs. In the erythromycin and penicillin strategies, drug costs dominate anaphylaxis and infection costs.

There is a trade-off between the two antibiotics. Penicillin is slightly less expensive whereas erythromycin is slightly more effective because it carries no risk of anaphylaxis. If the dentist is *absolutely certain* that the patient is not allergic to penicillin, then penicillin prophylaxis would be both less expensive and more effective than erythromycin and the marginal cost-effectiveness of penicillin compared to no prophylaxis would be $9,100 per QALY saved (analysis not shown).

Two published analyses have addressed the issue of endocarditis prophylaxis for dental procedures in patients with mitral valve prolapse. Bor and Himmelstein[12] concluded that no prophylaxis or prophylaxis with erythromycin is preferable to prophylaxis with a combination of parenteral and enteral penicillin because the number of deaths from penicillin reactions far outweighs the number of deaths from endocarditis that would be prevented by penicillin prophylaxis. Clemens and Ransohoff[15] found that the risks of parenteral penicillin outweigh the benefits; in addition, they concluded that oral penicillin prophylaxis spares life only in older adults with mitral valve prolapse, at a cost of $2.6 million per case of endocarditis prevented ($1.3 million per year of life saved).

Our analysis arrives at similar conclusions despite a major difference in the risk of infection. The other analyses estimated the risk of endocarditis to be 4.1–4.9 cases per million dental procedures.[12,15] In this analysis, the risk of septic arthritis was estimated to be 93 per million dental procedures. Their estimates were based on reported cases; our estimate was based on a cohort study.[16] If these infection rates are accurate, then we would expect to see more cases of endocarditis in patients with mitral valve prolapse after dental procedures than we see septic arthritis in patients with artificial joints only because mitral valve prolapse is much more prevalent than are prosthetic joints.

The derived probability of developing a joint infection rests on tenuous data and assumptions, including the probability of seeding a joint, the number of edentulous 65-year-olds, the frequency with which dentulous and edentulous patients see a dentist, etc. The sensitivity analysis of the probability of developing a joint infection (Table 2) encom-

TABLE 3—Per Capita Component Costs of Each Strategy

Component	Strategy		
	No Prophylaxis	Erythromycin	Penicillin
Antibiotics	$0.00	$3.73	$3.71
Anaphylaxis	$0.00	$0.00	$0.24
Joint infections	$1.49	$0.37	$0.15
Total	$1.49	$4.10	$4.09*

*Total differs from sum of components due to rounding

passes sensitivity analyses on each of those assumptions and probabilities. Central to this analysis, of course, is the assumption that a transient bacteremia during a routine dental procedure *can* cause an infection of prosthetic joint. Ideally, we could answer this question with a clinical trial or case-control study. Until that time, we rely on assumptions and previous studies and conclude that, from a patient's perspective, the benefits of taking prophylactic erythromycin prior to dental procedures outweigh the risks. Moreover, the cost to a payor per quality-adjusted year of life saved is relatively low, being comparable in a cost-effectiveness to estrogen therapy in the menopause,[17] treatment of mild hypertension,[18] and dietary treatment of hypercholesterolemia[19] (analyses done from a societal perspective).

APPENDIX

Probability of Developing an Artificial Joint Infection from a Transient Bacteremia During a Dental Manipulation

Ahlberg and colleagues[16] found that 0.29 per cent of patients with artificial joints developed hematogenous infections over an eight-year period (average length of follow-up not given). But only a portion of hematogenous infections, 10 per cent in Little's estimate,[2] are caused by transient bacteremias from dental manipulations. Knowing these values, the number of dental visits per year in the 65-year-old age cohort (1.3 visits per person per year),[20] and the percentage of these patients who do not wear dentures (and thus are at risk for seeding the joint), one can estimate a probability of a joint infection in the following manner:

Assuming that the *average* length of follow-up in Ahlberg's series was four years, we divide 0.0029 by 4 to get a rate of 7.3×10^{-4} hematogenous infections per patient per year. Because each patient, on average, sees the dentist 1.3 times each year, the rate of hematogenous infections per visit is 5.6×10^{-4}. Assuming only 10 per cent of hematogenous infections are caused by transient bacteremias from dental manipulations, we deduce the probability of developing a joint infection from a transient bacteremia in the dentist's chair to be 5.6×10^{-5}. This is the cumulative population risk; but this risk is borne only by the 60 per cent of patients who are dentulous. Thus the risk among dentulous patients is 9.3×10^{-5} or 9.3 per 100,000 dental procedures. Note that this risk may well be a lower bound because some of the patients in Ahlberg's series may have received prophylactic antibiotics and because his criteria for joint infection may have omitted indolent cases.

Short-term Morbidity from Taking Penicillin

Approximately 5 per cent of patients taking penicillin develop a hypersensitivity reaction, most commonly a rash.[21] We assumed that rash would last 48 hours, during which quality of life would be 95 per cent of baseline quality. To calculate the average disutility per patient, we multiplied the fraction of patients affected (0.05) by the product of the duration (48 hours) and disutility (1–0.95); the result is not if 7.5 minutes, which we deducted from the patient's life expectancy.

Short-term Morbidity from Taking Erythromycin

Here, we assumed that approximately 75 per cent of patients would suffer gastrointestinal side effects[22] lasting three to four hours, and that their quality of life during that time would be 95 per cent of baseline. The net average disutility is again 7.5 minutes per patient.

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Physical Activity and Risk of Cancer in the NHANES I Population

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Abstract: We studied the relation between self-reported physical activity and cancer in the first National Health and Nutrition Examination Survey (NHANES I) cohort, originally examined between 1971–75, and followed prospectively through the Epidemiologic Follow-up Study (NHEFS), conducted between 1982–84. Among 5,138 men and 7,407 women 25–74 years old, for nonrecreational activity we observed increased risk of cancer among inactive individuals compared to very active persons (for men, relative risk [RR] 1.8, 95% confidence interval [CI] = 1.4, 2.4; for women RR 1.3, 95% CI = 1.0, 1.8). These findings were unchanged after adjustment for cigarette smoking, body mass index (BMI), and other potential confounders. Sites which demonstrated stronger inactivity-cancer associations included colorectum (RR 1.6,95% CI = 0.7, 3.5) and lung (RR 1.6; 95% CI = 1.2, 3.5) among men, and breast (postmenopausal) (RR 1.7; 95% CI = 0.8, 2.9) and cervix (RR 5.2; 95% CI = 1.4, 14.5) among women, although these findings for women were based on relatively few cases. The association between inactivity and cancer was greater among persons of moderate (or lower) BMI, those cases occurring three or more years after baseline, and, in women, those more than 60 years old. In contrast, recreational exercise showed little relation to cancer, with the exception of prostate cancer. The results suggest that inactive individuals are at increased risk of cancer. (*Am J Public Health* 1989; 79:744–750.)

Introduction

The beneficial effects of physical activity on cardiovascular disease risk have been well documented.[1] Persons with other chronic illnesses such as osteoporosis and diabetes mellitus have likewise shown improved clinical status while on exercise regimens.[2,3] To a lesser degree, physical activity has also been studied with respect to risk of cancer. Early studies provided evidence for an inactivity-cancer relation,[4–6] but there have been conflicting data.[7,8] More recently, a population-based study of men[9] demonstrated an inverse association between incidence of colon cancer and level of occupational physical activity, which was corroborated by two subsequent reports.[10,11] Although these studies, and one of female cancers among former college athletes,[12] support the notion that greater physical activity is associated with reduced cancer risk, they have been limited to group estimates of physical activity based upon occupational classification or athletic status.

We examined the relation between self-reported activity level and the subsequent development of cancer in a large cohort of US men and women. The present investigation extends previous research by assessing the role of both recreational and nonrecreational physical activity, by evaluating relative body weight and other potential confounders of the activity-cancer association, and by including men and women.

Methods

We analyzed data available on public use tapes from the first US National Health and Nutrition Examination Survey (NHANES I) and the NHANES I Epidemiologic Follow-up Study (NHEFS). Details of their design have been presented elsewhere.[13,14] Between 1971 and 1975, 14,407 men and women ages 25–74 participated in NHANES I which was conducted by the National Center for Health Statistics (NCHS) and was designed to characterize a variety of health and nutrition-related parameters in a representative sample of the U.S. civilian population. Of the participants, 12,554 (87 per cent) were successfully traced and reinterviewed (14 per cent by proxy) during the period 1982–84 by NCHS as part of the NHEFS. All persons for whom baseline data existed concerning age, physical activity, stature and body weight, and who were without cancer at study entry (35 prevalent cases were identified through the baseline medical history interview and hospital records) were included in this investigation. Thus, the analytic cohort was comprised of 5,141 men (41 per cent) and 7,413 women (59 per cent), of whom 84 per cent were White. The median age at study entry was 48 years for women and 52 years for men.

Incident cases of cancer (excluding nonmelanoma skin cancer), confirmed through hospital record or death certificate review, served as endpoints for this study. Pathology reports were obtained for all cases. During the follow-up period (median length 10 years), 460 cancers developed among men (including 114 lung, 95 prostate, and 62 colorectum), and 399 among women (122 breast and 67 colorectum). Because relatively few cases (i.e., <35) occurred for specific cancer sites other than lung, prostate and colorectum in men, and breast and colorectum in women, they were not initially analyzed separately. Similarly, cancer of the colon and rectum were left combined because of very small case numbers (for men and women, respectively: colon, 36 and 48; rectum, 26 and 18), and the lack of any material difference between their respective associations with activity.

Two questions concerning physical activity were asked during the baseline evaluation. The first, dealing with nonrecreational activity, was: "In your usual day, aside from recreation, how active are you?" The possible responses were: "very active", "moderately active", or "quite inactive". The second question ("recreational exercise") asked: "Do you get much exercise in things you do for recreation?" "Much exercise", "moderate exercise", or "little or no exercise" were the available responses. Information concerning potential confounding risk factors was also obtained. Height and weight were measured at baseline using standard methods,[15] and the body mass index (BMI, weight (kg)/height $(m)^2$) was used as a measure of relative body fatness. Race was classified as White, Black, or other at baseline; because there were relatively few persons in the latter two categories, they were combined (i.e., non-Whites) for the present investigation. Maximum education attained was queried in the

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baseline interview, and relative economic status was estimated through the use of the NHANES I poverty index ratio, a measure of adjusted family income which compares a person's income, residence, and family size to the poverty level. Cigarette smoking status (never, former, current) and total pack-years were derived from either the baseline interview, or at follow-up interview. Pack-years was the product of the duration of smoking (in years) and the usual number of cigarettes smoked daily (divided by 20). Reproductive and gynecologic history (age of menarche, menopausal status, parity, and age at first live birth) and family history of breast cancer came from both baseline and follow-up interviews. Self-reported employment status, general health ("excellent" to "poor"), and history of a recent hospitalization (during the past year) were obtained at baseline, as was dietary energy and total fat intake information derived from a 24-hour recall questionnaire.

Relative risk estimates were computed using the proportional hazards model of SAS (PHGLM) in which incident cases of cancer served as study events.[16,17] Indicator variables were used for the two lower physical activity levels, leaving persons in the highest activity category as the reference group for both recreational and nonrecreational activity analyses. Age was included in all models as an independent, continuous variable. As indicated in the text, some analyses included adjustment (using indicator variables) for cigarette smoking status and pack-year history, economic status (using the poverty index ratio), reproductive and family breast cancer history, BMI, or dietary fat or energy intake, and for these analyses persons with missing data were excluded. In order to assess the potential impact of early stage, undiagnosed cancer on the physical activity-cancer relationship, we conducted stratified analyses based on length of follow-up i.e., <3 years versus ≥ 3 years. For the later follow-up period, cancers occurring within three years of study entry were excluded. Because none of the factors was found to be a significant confounder in these data, and covariate information was not available for all cases and noncases, only age-adjusted relative risks are presented in the Tables. Trend tests were based on the significance of a model trend variable for activity (scored 0,1,2).

TABLE 1—Age-adjusted[a] Mean (or distribution of) Baseline Characteristics According to Level and Category of Physical Activity by Sex

Nonrecreational Activity	Very Active	Moderately Active	Quite Inactive
Males	(2267)[b]	(2291)	(580)
Age (years)	49.7	54.0	55.4
% Current smokers	28.7	32.4	25.5
% Non-White	14.1	13.7	15.5
% Education >HS	17.7	30.3	31.9
Wt/ht^2 (kg/m^2)	25.6	25.8	25.6
Height (cm)	173.8	173.9	173.5
Energy intake (kcal)	2281	2077	1943
Females	(3102)	(3558)	(747)
Age (years)	45.4	49.4	52.6
% Current smokers	12.2	12.8	14.4
% Non-White	15.5	15.4	22.5
% Education >HS	19.6	23.1	16.8
Wt/ht^2 (kg/m^2)	25.1	25.8	27.3
Height (cm)	161.2	161.3	160.3
Energy intake (kcal)	1485	1463	1410
Recreational Exercise	**Much Exercise**	**Moderate Exercise**	**Little or no Exercise**
Males	(1191)	(2024)	(1922)
Age (years)	48.6	51.7	55.1
% Current smokers	29.9	32.5	27.4
% Non-White	9.8	11.2	19.8
% Education >HS	27.3	27.2	21.1
Wt/ht^2 (kg/m^2)	25.4	25.8	25.8
Height (cm)	174.3	174.0	173.3
Energy intake (kcal)	2328	2148	2067
Females	(1027)	(2736)	(3645)
Age (years)	45.6	46.5	49.9
% Current smokers	15.1	13.6	11.4
% Non-White	9.2	10.9	22.0
% Education >HS	27.8	24.9	16.2
Wt/ht^2 (kg/m^2)	24.4	25.0	26.5
Height (cm)	161.6	161.3	161.0
Energy intake (kcal)	1567	1453	1450

[a]Except mean age.
[b]Number of participants for activity categories in parentheses. For some subjects, information concerning smoking status, education, or energy intake was missing.

Results

Baseline characteristics of the study participants according to type and level of activity appear in Table 1. Most of the men and women participating in the study reported being either very active or moderately active in their nonrecreational activity, with only approximately 10 per cent being inactive. In contrast, over one-third of the men and nearly half of the women received little or no exercise in their recreational activities. Individuals reporting more daily activity or exercise tended to be younger and of lower BMI (the latter, primarily in women), while the least active persons were in general more likely to be other than White, less educated, and somewhat shorter in stature. One interesting exception concerns the smaller proportion of nonrecreationally very active men (17.7 per cent) with any education beyond high school compared with less active men. Energy intake was directly related to both types of activity, particularly among men, who also exhibited a much greater range in kilocalorie intake across activity categories.

Age-adjusted relative risk of cancer (all sites) was inversely related to nonrecreational physical activity level among men, demonstrating a nearly two-fold risk increase in the inactive group (Table 2). The increase in risk was smaller among women. Among men, the inverse association with nonrecreational activity was strongest for lung cancer, although it was also suggested for colorectal and prostate cancer. There was no similar association for either colorectal or breast cancer in women. Among inactive women, when breast cancer was analyzed according to menopausal status, an inverse association was seen in postmenopausal women [relative risk (CI) among the least active of 1.5 (0.7, 2.8)], while a direct association was suggested for premenopausal women [relative risk (CI) of 0.4 (0.1, 1.8) based on 46 cases]. In addition, there was an increase in risk for cancer of the cervix (n = 20 cases) by decreasing activity levels [relative risk (CI) = 1.0, 2.3 (0.7, 7.2), and 5.2 (1.4, 14.5)].

The findings concerning recreational activity were generally less impressive than for nonrecreational activity (Table 2). The risk of prostate cancer increased with decreasing levels of exercise, and there was a small, nonsignificant increase in risk of colorectal cancer among women receiving moderate or less exercise. Similar to the findings for nonrecreational activity, there was a suggestion of increased risk of breast cancer among postmenopausal women in the low exercise group [relative risk (CI) = 1.7 (0.8, 2.9)], with an

TABLE 2—Relative Risk[a] of Cancer According to Level and Category of Physical Activity by Sex

Nonrecreational Activity	# Cases	Very Active[b]	Moderately Active	Quite Inactive	Test for Trend p-value
Males					
All Sites	460	1.0	1.1 (0.9, 1.4)*	1.8 (1.4, 2.4)	0.0002
Colorectum	62	1.0	1.2 (0.7, 2.1)	1.6 (0.7, 3.5)	0.24
Lung	114	1.0	1.3 (0.9, 2.0)	2.0 (1.2, 3.5)	0.02
Prostate	95	1.0	0.9 (0.6, 1.3)	1.3 (0.7, 2.4)	0.78
Females					
All sites	399	1.0	1.0 (0.8, 1.3)	1.3 (1.0, 1.8)	0.16
Colorectum	66	1.0	1.3 (0.8, 2.2)	0.7 (0.3, 2.0)	0.98
Breast	122	1.0	0.9 (0.6, 1.3)	1.1 (0.6, 2.0)	0.92
Recreational Exercise	**# Cases**	**Much Exercise[b]**	**Moderate Exercise**	**Little or no Exercise**	
Males					
All sites	460	1.0	1.1 (0.9, 1.4)	1.2 (1.0, 1.6)	0.08
Colorectum	62	1.0	1.0 (0.5, 1.9)	1.0 (0.5, 1.9)	0.93
Lung	114	1.0	1.0 (0.6, 1.6)	0.9 (0.6, 1.5)	0.80
Prostate	95	1.0	1.2 (0.6, 2.2)	1.8 (1.0, 3.3)	0.02
Females					
All sites	399	1.0	1.0 (0.7, 1.3)	1.0 (0.7, 1.3)	0.82
Colorectum	66	1.0	1.2 (0.5, 2.7)	1.2 (0.6, 2.8)	0.63
Breast	122	1.0	0.9 (0.5, 1.6)	1.0 (0.6, 1.6)	0.98

[a]Relative risk estimated from the proportional hazards regression models which include age as an independent variable.
[b]Reference categories.
*95% confidence interval in parentheses.

TABLE 3—Relative Risk[a] of Cancer (all-sites) According to Level of Nonrecreational Activity and Recreational Exercise by Sex

	Males				Females			
Recreational Exercise	# Cases	Very Active	Moderately Active	Quite Inactive	# Cases	Very Active	Moderately Active	Quite Inactive
Much	88	1.0[b]	1.1 (0.7, 1.9)*	1.7 (0.4, 7.0)	53	1.0[b]	1.0 (0.5, 2.1)	1.3 (0.2, 9.3)
Moderate	168	1.0 (0.7, 1.4)	1.1 (0.8, 1.5)	1.4 (0.7, 2.9)	137	1.0 (0.7, 1.5)	1.0 (0.7, 1.4)	1.9 (0.8, 4.5)
Little or None	204	1.0 (0.7, 1.5)	1.2 (0.8, 1.6)	1.9 (1.3, 2.7)	209	0.8 (0.6, 1.3)	1.0 (0.7, 1.4)	1.2 (0.8, 1.8)

[a]Relative risk estimated from the proportional hazards regression models which include age as an independent variable.
[b]Reference category (very active and much exercise).
*95% confidence interval in parentheses.

opposite association premenopausally [relative risk (CI) = 0.6 (0.3, 1.2)].

Adjustment for potential confounding risk factors [race, economic and smoking status, BMI, and energy intake (all sites); BMI and dietary fat intake (colorectum); smoking status and cigarette pack-year history (lung); age at menarche and menopause, parity, age at first birth, as well as family history of breast cancer, BMI (or stature), and dietary fat intake (breast cancer); economic status (cervix)] did not materially alter any of the findings concerning either activity type.* For example, the multivariate adjusted relative risk of all sites cancer remained unchanged: 1.0, 1.1, 1.8 for the decreasing levels of nonrecreational activity in men; 1.0, 1.1, 1.2 for women. Similarly, the results were not affected by adjustment for employment status (working versus not) or general health status (subjective report of health or a hospitalization in the past year).

Table 3 shows that compared to persons who claimed to be both very active and exercised, it was generally only those reporting being less active who were at increased cancer risk. (Cancer incidence (per 100,000 person-years) for the reference category was 817 for men and 560 for women.) Joint classification for nonrecreational activity and exercise therefore yielded relative risk estimates comparable to those for nonrecreational activity alone, without evidence of synergism.

Because many of the above results implicated nonrecreational inactivity specifically as a cancer risk factor, stratified analyses for age, BMI, and follow-up time (factors possibly related to inactivity) were conducted and are shown in Tables 4 and 5. In addition, Whites and all others were analyzed separately (Table 4). The risk increase associated with diminished activity was observed in both younger (25–59 years of age) and older men regardless of race (Table 4).

*Data available on request to author.

TABLE 4—Relative Risk[a] of Cancer (all-sites) According to Level of Nonrecreational Activity by Sex and Subgroups of Age, Race, Body Mass Index, and Follow-up Time

	Males					Females				
	# Cases	Very Active[b]	Moderately Active	Quite Inactive	Trend Test p-value	# Cases	Very Active[b]	Moderately Active	Quite Inactive	Trend Test p-value
Age										
≤60 years	119	1.0	1.5 (1.0, 2.3)*	1.6 (1.1, 2.2)	0.004	159	1.0	1.2 (0.7, 1.9)	0.9 (0.7, 1.3)	0.75
>60 years	341	1.0	1.1 (0.8, 1.3)	1.9 (1.3, 2.4)	0.007	240	1.0	1.0 (0.8, 1.3)	1.5 (1.0, 2.3)	0.11
Race										
White	388	1.0	1.0 (0.8, 1.3)	1.8 (1.4, 2.5)	0.002	346	1.0	1.0 (0.8, 1.2)	1.2 (0.9, 1.8)	0.47
non-White	72	1.0	1.7 (1.0, 2.9)	1.8 (0.9, 3.8)	0.05	53	1.0	1.7 (0.9, 3.4)	2.1 (1.0, 4.8)	0.05
Body Mass Index										
low (<22.0)	107	1.0	1.4 (0.9, 2.2)	3.1 (1.9, 5.2)	<0.0001	92	1.0	1.1 (0.7, 1.7)	1.8 (0.9, 3.4)	0.19
moderate (22–26)	163	1.0	1.5 (1.1, 2.2)	1.9 (1.1, 3.0)	0.005	124	1.0	1.1 (0.8, 1.7)	2.0 (1.2, 3.5)	0.03
high (>26.0)	190	1.0	0.8 (0.6, 1.1)	1.2 (0.7, 1.3)	0.83	183	1.0	0.9 (0.7, 1.3)	0.9 (0.5, 1.4)	0.57
Follow-up time										
< 3 years	122	1.0	0.9 (0.6, 1.4)	1.4 (0.8, 2.4)	0.38	114	1.0	0.9 (0.6, 1.4)	1.3 (0.8, 2.4)	0.61
≥ 3 years	338	1.0	1.2 (0.9, 1.5)	2.0 (1.4, 2.8)	0.0002	285	1.0	1.1 (0.8, 1.4)	1.3 (0.9, 1.9)	0.21

[a]Relative risk estimated from the proportional hazards regression models which include age as an independent variable.
[b]Reference categories.
*95% confidence interval in parentheses.

The association was restricted, however, to older women who were not White, and to low and moderate relative weight men and women, with a three-fold risk increase among inactive lean men compared to very active lean men. While length of follow-up did not affect the inactivity-cancer association in women, a somewhat stronger inverse relation was seen among men for cancers occurring three or more years after baseline.

Additional evidence for effect modification by age and BMI appears in Table 5. There was an inverse cancer-activity association only among younger and older men who were of average BMI or less, and in younger, heavier men. Among women, an effect was observed only for the older and leaner category.

The combined effects of race and either age or BMI were also evaluated. Among men, a somewhat stronger association between inactivity and cancer was observed in younger (versus older) Whites, and in other than older White women (Table 6). In both race groups, the association was essentially restricted to those in the lower BMI category. Adjustment for cigarette smoking did not alter the latter finding. Table 7 shows that the inverse activity-cancer association was evident primarily among non-White women, of either age group, and women of lower BMI.

TABLE 5—Relative Risk of Cancer (all-sites)[a] According to Nonrecreational Activity Level Stratified by Age, Body Mass Index, and Sex

Age (yrs)	Body Mass Index (kg/m²)	# Cases	Very Active[b]	Moderately Active	Quite Inactive	Test for Trend p-value
Males						
≤60	≤26	63	1.0	2.2 (1.3, 3.8)*	2.0 (0.9, 4.6)	0.01
	>26	56	1.0	1.0 (0.6, 1.8)	2.2 (1.0, 4.8)	0.17
>60	≤26	207	1.0	1.3 (1.0, 1.8)	2.5 (1.7, 3.7)	<0.0001
	>26	134	1.0	0.7 (0.5, 1.0)	0.9 (0.5, 1.6)	0.30
Females						
≤60	≤26	101	1.0	1.2 (0.8, 1.8)	0.9 (0.4, 2.2)	0.72
	>26	58	1.0	1.2 (0.7, 2.1)	1.0 (0.4, 2.5)	0.76
>60	≤26	115	1.0	1.2 (0.8, 1.8)	2.8 (1.7, 4.8)	0.001
	>26	125	1.0	0.8 (0.6, 1.2)	0.8 (0.5, 1.4)	0.37

[a]Relative risk estimated from the proportional hazards regression models which include age as an independent variable.
[b]Reference category.
*95% confidence interval in parentheses.

Discussion

Most previous investigations of this area have examined either occupational history or involvement in athletics. Several studies demonstrate an inverse association between occupational physical activity, based on job classification, and the development of malignancy.[4,6,9–11] Siversten and Dahlstrom,[4] in a historically interesting report, showed that the death rate from carcinoma was higher among the unemployed, and inversely related to estimated occupational "muscular activity." The authors attributed their observations to metabolic products of deficient muscular activity which altered body fluids and, therefore, adult epithelial tissue. The effect of age and the tendency for ill individuals to be unemployed or reduce their activity were not, however, adequately addressed. Among railroad workers, Taylor[6] demonstrated approximately one-third lower cancer death rates for "section men" (those involved in heavy manual

TABLE 6—Relative Risk[a] of Cancer (all-sites) According to Level of Nonrecreational Activity Stratified by Race, Age, and Body Mass Index, for Males

	White					All Others				
	# Cases	Very Active[b]	Moderately Active	Quite Inactive	Trend Test p-value	# Cases	Very Active[b]	Moderately Active	Quite Inactive	Trend Test p-value
Age (years)										
≤60	98	1.0	1.4 (0.9, 2.1)*	2.2 (1.2, 4.0)	0.008	21	1.0	2.3 (0.9, 5.9)	1.0 (0.1, 7.7)	0.28
>60	290	1.0	1.0 (0.7, 1.2)	1.7 (1.2, 2.4)	0.03	51	1.0	1.5 (0.8, 2.8)	1.9 (0.8, 4.4)	0.10
Body mass index (kg/m^2)										
≤26	223	1.0	1.3 (1.0, 1.8)	2.3 (1.6, 3.3)	0.0001	47	1.0	2.8 (1.3, 6.0)	3.6 (1.4, 9.3)	0.003
>26	165	1.0	0.8 (0.6, 1.1)	1.3 (0.8, 2.2)	0.91	25	1.0	0.8 (0.4, 1.9)	0.5 (0.1, 2.4)	0.38

[a]Relative risk estimated from the proportional hazards regression models which include age as an independent variable.
[b]Reference category.
*95% confidence interval in parentheses.

TABLE 7—Relative Risk[a] of Cancer (all-sites) According to Level of Nonrecreational Activity Stratified by Race, Age, and Body Mass Index, for Females

	White					All Others				
	# Cases	Very Active[b]	Moderately Active	Quite Inactive	Trend Test p-value	# Cases	Very Active[b]	Moderately Active	Quite Inactive	Trend Test p-value
Age (years)										
≤60	139	1.0	1.1 (0.8, 1.6)*	0.8 (0.4, 1.6)	0.90	20	1.0	1.9 (0.7, 5.3)	2.1 (0.5, 8.6)	0.19
>60	207	1.0	0.9 (0.7, 1.3)	1.4 (1.0, 2.2)	0.24	33	1.0	1.6 (0.7, 3.8)	2.0 (0.7, 5.6)	0.17
Body mass index (kg/m^2)										
≤26	190	1.0	1.1 (0.8, 1.4)	1.6 (1.0, 2.6)	0.12	26	1.0	1.9 (0.7, 5.1)	4.6 (1.6, 13.6)	0.006
>26	156	1.0	0.8 (0.6, 1.2)	0.9 (0.5, 1.5)	0.50	27	1.0	1.6 (0.7, 4.0)	0.9 (0.2, 3.5)	0.85

[a]Relative risk estimated from the proportional hazards regression models which include age as an independent variable.
[b]Reference categories.
*95% confidence interval in parentheses.

labor) compared to less active clerks and switchmen. In three more recent studies,[9–11] the risk of colon cancer was increased by between 30 and 100 per cent among men employed in sedentary occupations. Garabrant, *et al*,[9] suggested that reduced stimulation of colonic peristalsis among those least active could account for the association.

All of the above investigations involved working men, and utilized estimated physical activity based on occupational categories. Our demonstration of a nearly two-fold risk increase among inactive men relative to those more active is consistent with these previous studies, both in terms of the magnitude of the association and the specific implication of inactivity or "sedentariness". The present findings concerning nonrecreational activity are also consistent with the specific association most recently reported[9–11] between colon cancer in men and occupational inactivity. In our study, however, the relation exists for cancer of the rectum as well, but was not observed in women. Nonrecreational activity was positively related to employment in the present data: 79 per cent of the "very active" men were employed compared with only 67 and 51 per cent of the moderately active and inactive men, respectively (49, 39, and 35 per cent for women in the same activity categories). Conversely, only 7 per cent of men with jobs reported being inactive, compared with 19 per cent of the unemployed men. These apparent similarities notwithstanding, our study demonstrated an activity-cancer association which is independent of employment status, since adjustment for this factor did not affect our results, and the nonrecreational activity association was observed in both the employed and unemployed (data not presented). Therefore, while inactive persons were less likely to be working, it was not their lack of work which increased their risk of cancer.

The study of athletes, another approach to investigating the relation between cancer and physical activity (or more specifically, fitness) has yielded conflicting data.[5,7,8,12] Two early studies[5,7] lacked control groups, and may have suffered both from biased samples and inadequate consideration of age. A retrospective cohort study based on death certificates and athletic status in college showed a small excess of age-stratified (by 10-year birth cohorts) cancer mortality, particularly for cancer of the prostate, among Harvard "letter men" (varsity athletes) compared to either athletes not having received a "letter", or to students who had applied for (but did not receive) a gymnasium locker.[8] The authors could not rule out other confounding effects such as smoking or body size, however. In a similar investigation, female college alumni who had actively participated in one or more team sports during college reported less subsequent cancer of the breast, uterus (including cervix), and ovary than did nonathlete alumni.[12] This association, which persisted after adjust-

ment for several breast cancer risk factors, was evident primarily in women 50–70 years old. However, the number of cancers was small (69 breast, 37 all others combined). One other recent report demonstrated a direct relation between heart rate and total cancer mortality that was unchanged after adjustment for several factors.[18] The authors interpret their findings in light of theories concerning adrenergic activity and psychological stress, although they do not exclude the effects of habitual exercise or sedentary life-style. Interestingly, two sites showing a strong, positive association in this study were lung and colon (a finding similar to ours) but only for the latter site did it remain significant after adjustment for smoking and serum cholesterol.

We demonstrated a positive association with recreational exercise only for prostate cancer, a finding opposite that observed among Harvard athletes.[8] The suggestion of decreased breast cancer risk among postmenopausal women receiving much exercise in our study is consistent with the study of female college athletes mentioned.[12] In that investigation, earlier athletic status provided some risk reduction in the postmenopausal age range only, with no demonstrable athlete-nonathlete difference premenopausally. We did observe increased risk for the high exercise category among premenopausal women. Also, data for cancer of the cervix and endometrium (not shown) are suggestive of a protective effect for a high level of exercise, similar to that observed previously for female athletes,[12] and not accounted for by race, economic status, or body mass index. Unfortunately, the number of cases does not permit a more detailed evaluation of these associations in women, which must therefore remain tentative until investigated by others. The latter associations notwithstanding, given the positive findings for nonrecreational activity, the lack of a stronger overall association for exercise in the present investigation is somewhat surprising. It is possible that the effects of recreational activity are diminished because the duration of daily exercise (and hence its potential influence on disease) is generally much shorter than either athletic training or work-related activity. Alternatively, the observed association for nonrecreational activity may be unrelated to physical fitness; the implication of inactivity *per se* could be interpreted as being supportive of this. It has already been pointed out that the two questions used in the NHANES surveys reflect different aspects of daily activity.[19] Interestingly, we did find that the relation between cancer and nonrecreational inactivity was suggested at all levels of reported daily exercise in men, and for the moderate and little or no exercise groups in women.

The implication of inactivity as a risk factor also suggested to us a possible role for characteristics or illnesses which would both reduce activity and predispose to the development of cancer. Although the appearance of a somewhat stronger inactivity-cancer association among leaner individuals and older women could be interpreted as supporting such a role, several of our findings argue against confounding by prior illness or infirmity. The activity association was observed among: 1) those followed for at least three years before developing cancer (in men, the association was even greater in this subgroup); 2) younger as well as older men; 3) the actively employed (excluding the unemployed and homemakers); 4) nonsmokers; and, possibly most importantly, 5) persons reporting good or excellent general health as well as those denying any hospitalization during the year preceding study entry (data not shown). Thus, it is unlikely that ill health could account for our findings.

The present investigation is unique in the collection of self-reported information concerning daily nonrecreational activity and physical exercise. Such data offer the advantage of individual estimates which may be more specific than job categories or athletic status decades earlier. While an effect from the subjective interpretation of the interview questions cannot be excluded, misclassification from this should only have caused our risk estimates to be reduced. It is noteworthy that we demonstrated increased cancer risk among inactive individuals in spite of this. One level of validation of the NHANES I activity classification is afforded by the caloric intake data which are available. Both age-adjusted mean energy intake (Table 1) and energy not used for resting metabolic rate (RMR) (energy intake minus predicted RMR,[20] a crude estimate of physical activity energy expenditure) increased with activity level. Furthermore, as already discussed, nonrecreational activity was positively related to employment status. Therefore, while direct validation of activity was not feasible in the present study, these indirect data support the value of the self-reports.

In contrast to (and an advantage over) most previous studies, the potential confounding effects of several relevant factors were also evaluated in the present investigation. We observed only relatively weak and inconsistent relations between activity level and most factors studied, and there was no evidence for confounding. Residual confounding due to factors we could not evaluate remains possible, however. It has been suggested that BMI has this potential, especially with respect to colon cancer.[21] While increased BMI is a risk factor for selected sites (including colon) in this cohort,** we observed little change to the relative risk estimates for activity groups upon adjustment for BMI. A stronger inactivity-cancer association was evident primarily among low or moderate BMI individuals for both men and women, particularly among non-Whites. Relative body mass may therefore modify either the effect or level of activity. That is, leanness may be necessary for activity to affect the development of cancer; for example, through an altered metabolic characteristic. Alternatively, it may only be leaner individuals who are active enough to achieve the protective influence, or who more accurately classify their activity level and afford a better test of the hypothesis.

Studies providing more detailed assessment of individual physical activity, and which include women, are needed. Research aimed at elucidating the relevant mechanism(s) would also be useful.

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International Conference on Emergency Health Care Development Invites Papers

The International Conference on Emergency Health Care Development, to be held August 15–18, 1989 at the Hyatt-Regency Hotel, Crystal City at National Airport, in the Washington, DC metropolitan area, invites interested persons to submit abstracts on work related to emergency and disaster health care development. Topics on organization, administration, and management will be considered. The conference is intended to serve as a forum and training ground for fostering the discussion of methods for improving emergency health care worldwide.

Attendees wishing to present papers or studies must admit an abstract on the form provided by the conference organizer. To obtain an abstract form, instructions, and a preliminary agenda, contact Medical Care Development International (MCDI), Conference Organizer, 1742 R Street, NW, Washington, DC 20009, telephone: 202/462-1920. Abstracts must be received by June 15, 1989 to be considered for the conference agenda. To speed receipt, prospective participants are encouraged to send a copy of the abstract by telefax in care of MCDI, 202/265-4078, at the same time the original is mailed.

The conference will focus on the development of emergency health care systems, both pre-hospital and in-hospital health care services that respond to sudden illness or injury. The objectives of the international conference are to:

- Clarify the linkages between emergency health care, local and national health care services, and national disaster management systems;
- Present concrete methods for developing or improving emergency health care and disaster response management capabilities within societies which differ widely in resources and characteristics;
- Demonstrate processes by which different emergency health care services have been developed; and
- Propose international emergency health care development goals for the next decade.

The conference agenda is designed to permit special interest tracks, including: Training/Education, Management, Primary Health Care, Disaster Preparedness/Response, and Injury Control/Accident Prevention.

Conference sponsors include: US Department of Health and Human Services, Public Health Service, Health Resources and Services Administration; US Department of Transportation, National Highway Traffic Safety Administration; and the World Health Organization/Pan American Health Organization; in collaboration with United Nations Children's Fund and United Nations Disaster Relief Office.

Ethnic Variation in Cholecystectomy Rates and Outcomes, Manitoba, Canada, 1972–84

MARSHA M. COHEN, MD, T. KUE YOUNG, MD, AND KAREN M. HAMMARSTRAND, MA

Abstract: We used population-based data from the Province of Manitoba's universal health insurance plan to compare the cholecystectomy experience of Native Americans and non-Natives from 1972 to 1984. The age-adjusted cholecystectomy rates for Native females were higher than for non-Native females with the peak rate occurring at age 30–39 for Native Americans and at age 60–69 for non-Natives. The rates for males were three times lower than for females and did not differ between Natives and non-Natives. Native Americans were more likely readmitted to hospital for surgical complications than non-Natives and this held true after controlling for age, sex, rural versus urban residence, teaching versus non-teaching hospital, multiple discharge diagnoses or complex versus simple cholecystectomy (relative odds 1.46, 95 per cent confidence interval 1.17, 1.18). The explanation for the relatively high rates of cholecystectomy among Native American females may be related to high rates of known risk factors for gallstone disease (such as obesity and high parity). However, the higher rates of surgical complications require further study. (*Am J Public Health* 1989; 79:751–755.)

Introduction

A high prevalence of gallbladder disease has been demonstrated in several North American Indian tribes, such as the Pima in Arizona,[1] the Chippewa in north Minnesota,[2] and the Micmac in Nova Scotia.[3] There is speculation that the rise in gallbladder disease among Native Americans constitutes a "New World Syndrome" of metabolic diseases which includes also obesity and diabetes, probably with a common genetic and environmental basis.[4]

While the occurrence of gallbladder disease among Native Americans has been well studied, the rates and outcomes of its surgical treatment among these groups have received limited attention.[5,6] In Manitoba, the existence of population-based data on all health service utilization under the province's universal health insurance plan provides an opportunity to examine the cholecystectomy experience of all residents since the 1970s, and to compare rates and outcomes between Native Americans and the rest of the population.

Methods

For residents of the Province of Manitoba (population about one million, about 50 per cent urban, 50 per cent rural), all physician and hospital care is covered by the Manitoba Health Services Commission (MHSC). Records of health care utilization are routinely filed with the MHSC and, since there are no premiums or user charges, the health insurance claims are known to be complete and highly reliable for the study of surgical rates.[7] Manitoba hospital claims correspond closely to the hospital medical record as to the date of surgery, major surgical procedures performed, and important diagnoses. Record linkage is possible using masked numerical identifiers since neither names nor addresses are released to the researchers. In addition, the MHSC also maintains a continuously updated population-based registry (census) of all residents eligible for inclusion in the plan.

For this study, all hospital records containing a procedure code for gallbladder surgery for the years 1972–78 (ICD-8 codes 43.0–43.9), and for 1979–84 (ICD-9 codes 50.1–51.99) were selected. Status Native Americans were identifiable from special residence codes. This code, however, did not include non-Status Native Americans or Metis. This determination of Indian status in Canada is legalistic and complex, based on the Indian Act. Metis is a distinct cultural group composed of descendants of mixed Native-White marriages during the early settlement of the Canadian West. The US practice of designating people of aboriginal descent with "blood quantum" is not used in Canada. However, because Natives constitute less than 5 per cent of the provincial population, the cholecystectomy rates for the rest of the population would be only slightly overestimated by including the non-Status groups. Data available for each case included demographic information, diagnoses associated with the admission (up to eight diagnoses are recorded), specialty of the surgeon (1972–78 only), and type of hospital (urban teaching, urban non-teaching, rural large, rural small).

From these computer files, we determined the number of cases by age, sex, type of hospital where surgery had taken place, specialty of the attending surgeon, and region of residence within the province (1979–84 only). Using population counts from the MHSC, we next determined the age-adjusted cholecystectomy rate (by the direct method with the 1981 total provincial population as the standard) and the age-specific rates for males, females, Native and non-Natives. The denominators for these rates were obtained from the MHSC registry. Many of the analyses were divided into two time periods 1972–78 and 1979–84 to correspond to the eighth and ninth revisions of the International Classification of Diseases in use at the time and to examine time trends.

Since deaths from cholecystectomy or gallstone disease are rare, (in Manitoba, 0.7 per cent up to two years after surgery[8]), we used readmissions after surgery as the measure of morbidity associated with cholecystectomy. This method involved computer linkage of the original surgical event with subsequent hospitalizations due to known complications associated with the original surgical procedure.[9] In brief, specialist physician panels judged all readmissions to hospital after cholecystectomy and determined which were related to the original surgery. Computer algorithms using diagnostic codes were then developed based upon the physicians' judgments. Agreement between the computer algorithms and the physicians' judgments was very high.[9] A list of the

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TABLE 1—Demographic and Medical Care Characteristics by Time Period and by Native Status

Demographics	1972–78		1979–84	
	Native American (n=631)	All Others (n=22,917)	Native American (n=696)	All Others (n=13,139)
Mean Age (SD)	39.8(15)	50.0(16)	39.4(15)	51.8(17)
% Female	88.0	72.0	82.0	68.6
% Hospital Type				
Urban teaching	28.7	25.7	26.5	19.5
Urban non-teaching	19.8	44.8	20.5	49.5
Rural 100+ beds	29.6	14.6	31.9	16.8
Rural <100 beds	21.9	14.8	11.9	14.1
% Specialty of Surgeon				
General surgeon	56.3	61.1	NA	NA
Other Surgeon	2.5	2.2	NA	NA
Rural general practice	34.2	18.4	NA	NA
Urban general practice	3.7	4.2	NA	NA
Other physician	3.3	14.1	NA	NA
% Residence				
Rural	NA	NA	30.7	29.9
Northern	NA	NA	42.7	1.5
Urban	NA	NA	26.6	68.6

NA = Not available

reasons for readmission up to 120 days which are included in the study is found in the Appendix.

In order to determine which factors might be associated with the risk of readmission, we performed a multiple logistic regression[10] which included age, sex, type of hospital, urban versus rural residence, and whether the original surgery was a simple or a complicated cholecystectomy; as a case-mix measure, we used the number of admission diagnoses associated with the original surgical admission.[11]

Results

Demographic and medical care characteristics of the cases by time period and Native status are seen in Table 1. In both periods the Native population was younger than the rest of the population and included proportionally more women. The majority of Native cases were operated upon in non-teaching hospitals, most of which were in rural areas in contrast to the rest of the population. Data on specialists were only available for the earlier time period. Native Americans were more likely to have received their operation from a rural general practitioner (34.2 per cent) than the rest of the population (18 per cent). In keeping with the distribution of the population of the province, the Native groups were much more likely to live in the north.

The total age-adjusted rate of cholecystectomy over time by sex and Native status is seen in Figure 1. Not surprisingly, the rate for females was about three times higher than that for males. For non-Native females, the rate of operation fell markedly over time. However for Native women, the rates remained high over the 13 years. For males, the rates for non-Natives fell slightly but rose somewhat for Natives. The age-specific rates for Native females peaked at ages 40–49 in 1972–78 and at ages 30–39 for 1979–84 (Figure 2) while for other females, the peak rates were at ages 60–69. The rates for Native females were higher than those for other females except among the elderly. For males, the highest age-specific rates were seen at ages 70–79 for both racial groups (Figure 3).

Table 2 shows the proportion of cases who had readmissions following surgery for the two time periods. Native Americans were more likely to be readmitted to hospital and were more likely to have multiple admissions. In the second time period, Native Americans were still more likely to be

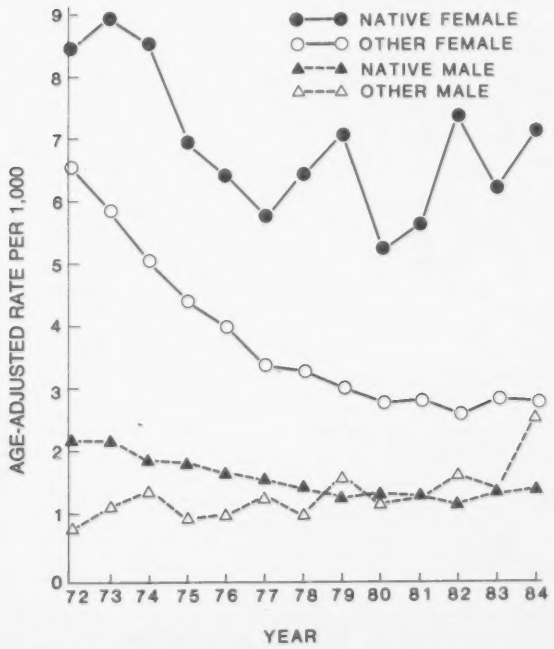

FIGURE 1—Age-adjusted Cholecystectomy Rate per 1,000 Population 1972–84 by Native Status and Gender, Manitoba, Canada

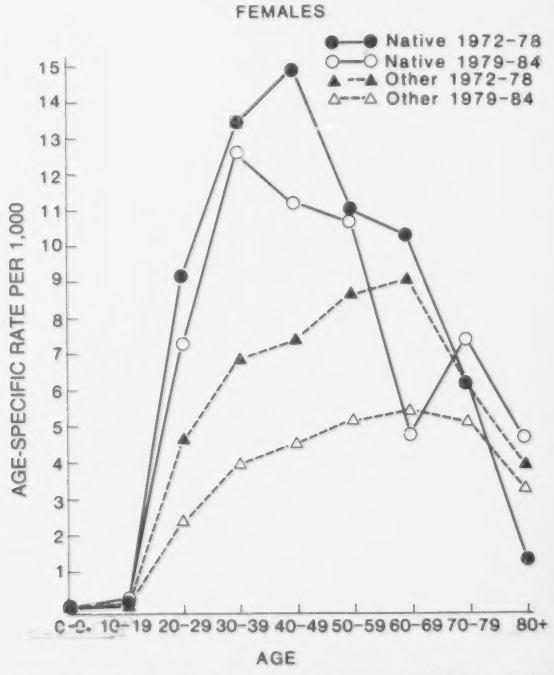

FIGURE 2—Age-specific Cholecystectomy Rate per 1,000 Females, by Native Status and Time Period (1972–78, 1979–84), Manitoba, Canada

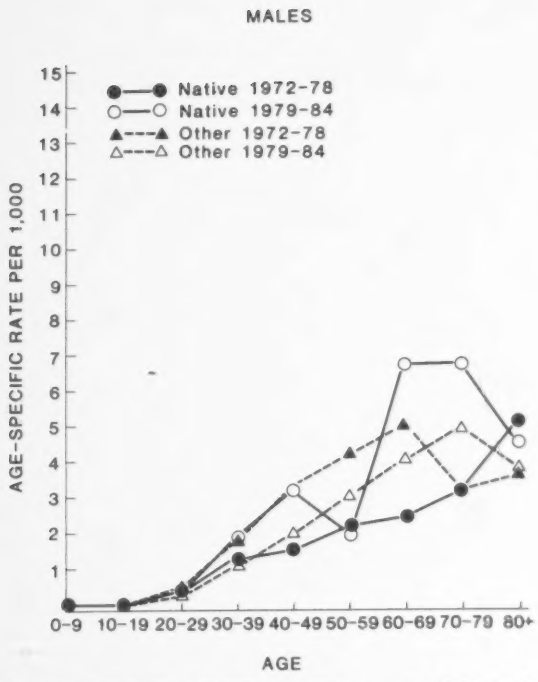

FIGURE 3—Age-specific Cholecystectomy Rate per 1,000 Males, by Native Status and Time Periods (1972–78, 1979–84), Manitoba, Canada.

readmitted than the rest of the population but the proportion of Natives having a readmission fell from 11.9 per cent to 7.8 per cent.

The results of the multiple logistic regression are seen in Table 3. Advancing age was associated with an increased risk of readmission to hospital for a surgically related complication. Females were less likely than males to be readmitted. Patients living in rural areas were more likely to be readmitted to hospital and having had the original surgery in a rural hospital (large or small) or at a Winnipeg teaching hospital placed one at elevated risk as compared to a nonteaching urban hospital. For those with multiple diagnoses (at time of the original surgery), there was a risk of 1.72 for readmission as compared to those whose only diagnosis was cholelithiasis. Complex surgery was not a risk factor for readmission as compared to simple cholecystectomy. After adjustment for all factors in Table 3, being a Native American placed one at 1.46 odds for readmission (95 per cent confidence interval = 1.17, 1.81).

Discussion

This study demonstrates that Native American women in Manitoba experienced a high rate of cholecystectomy. The age of operation was much earlier; they suffered from a high rate of postoperative readmissions; and the rate of this surgery has not declined much over time. Two explanations, not necessarily mutually exclusive, can be offered: Native women are at higher risk for developing gallbladder disease; Native Americans receive differential care within the health care system of Manitoba.

The high risk for gallbladder disease among North American Native Americans has been recognized for some time, particularly among tribes in the southwestern USA.[2,3,5,6,12,13,14] All Native groups studied had higher prevalence than the White population as exemplified by the Framingham Study,[15] with the Pima occupying the extreme high end of the spectrum, similar to their position with regard to diabetes.

A female excess is observed among both Natives and the rest of the population, although the female/male ratio is higher among the former. Cholecystectomy rates are not synonymous with prevalence of gallstones and may be affected by such factors as lack of agreement about indications for surgery.[16] However, virtually every study—whether autopsy, population-based, or of surgical rates—demonstrates a female/male ratio of at least 2:1.[17] Therefore the female excess in cholecystectomy rates likely reflects an excess in gallstone disease in Manitoba women. It can be speculated that the greater female excess and younger age of onset among Natives can be attributed to the higher prevalence of obesity and greater fertility among younger Native women. Both obesity and parity have been implicated as risk factors for gallbladder disease,[17–19] although the evidence is less consistent for Natives.[1–3,14,15] According to the 1981 Canada Census, an average of 4.8 children were born to ever-married Native women living on reserves, compared to 2.5 among others. Thirty-five per cent of Native women gave birth to six or more children, compared to only 8 per cent among non-Natives.[20] In terms of obesity, the Nutrition Canada Survey during the early 1970s indicated that Native women had higher mean weight than Canadians nationally at all age groups, whereas in men higher weights were observed only among those over the age of 50.[21]

Although the health care utilization pattern among Natives differs from that of the rest of the population, this is not sufficient to account for the higher surgical and readmission rate. The persistently high female rate is likely the result of a sustained disease burden unaffected by changing surgical practice in the province. The reason for the higher readmission rate is not known. Natives were as likely to receive bile duct exploration and the overall length of stay was also similar. Thus the differences were not likely due to retained stones at initial surgery.

Ethnic differences in post-cholecystectomy complications persist after controlling for various demographic and

TABLE 2—Percentage Readmitted to Hospital after Cholecystectomy by Time Period and Native Status

	1972–78			1979–84		
	Native American	All Others	Rate Ratio (95% CI)	Native American	All Others	Rate Ratio (95% CI)
First readmission	9.0	3.8	2.4(1.8, 3.1)	6.0	4.1	1.5(1.1, 2.0)
Second readmission	2.2	0.4	5.0(2.9, 8.7)	1.4	0.7	2.0(1.1, 3.9)
Third readmission	0.6	0.1	5.8(2.1, 15.9)	0.3	0.2	1.5(0.3, 4.5)
All readmissions	11.9	4.3	2.7(2.2, 3.5)	7.8	5.0	1.6(1.2, 2.0)

TABLE 3—Relative Odds for Readmission to Hospital after Cholecystectomy (logistic regression results)

Variables	Reference	Relative Odds	95% Confidence Interval
Age	60+/<40	1.88	1.64, 2.15
	40–60/<40	1.17	1.02, 1.35
Sex	Females/males	0.77	0.69, 0.86
Residence	Rural/urban	1.17	1.03, 1.30
Hospital type	Rural large/urban NT*	1.50	1.32, 1.74
	Rural small/urban NT*	1.36	1.14, 1.63
	Urban teaching/urban NT*	1.29	1.10, 1.45
Discharge diagnosis	Multiple diagnosis/ cholelithiasis only	1.72	1.53, 1.93
Type of gallbladder surgery	Cholecystectomy plus other procedure/cholecystectomy only	1.01	0.94, 1.12
Native status	Native/Other	1.46	1.17, 1.81

*Urban non-teaching hospital

medical care factors in a multivariate analysis. However, the types of "risk factor" information routinely available from health insurance claims data is limited. Differences in the distribution of lifestyle (smoking, diet, alcohol, physical activity), socioeconomic status, anthropometric and physiological indices (obesity, lipids, blood sugar, etc.), and parity may account for the observed "ethnic" differences and the continuing high rates.

Most studies of post-cholecystectomy complications tend to derive from one or more teaching hospitals or from one surgeon's experience.[22,23] These studies give complication rates reflecting only immediate postsurgical morbidity[24] and underestimate the true rate of adverse events which may occur after discharge from hospital.[8] Our study is population-based and includes all readmissions even if they occurred at different hospitals than the original surgical event. While not a definitive measure of complications of cholecystectomy, the use of readmissions data served to illustrate the differences between the Native experience and that of the rest of the population.

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(continued)

APPENDIX
Diagnoses, ICDA-8 CODES, ICD-9-CM Codes Used to Study Complications of Surgery at Readmission

Diagnoses/Complications	ICDA-8 Codes	ICD-9-CM Codes
General Complications		
Postoperative hemorrhage	998.1, 629.5	998.1, 623.8
Anemia	280, 285	280, 285
Postoperative wound infection, peritonitis, cellulitis	998.5, 567, 682.9	998.5, 567, 682.9
Pelvic inflammatory disease	616.0	614
Pulmonary embolism and infarction	450	415.1
Phlebitis and thrombophlebitis	451	451
Other venous embolism and thrombosis	453	453
Cystitis, pyelonephritis urinary tract infection	595, 590.1, 599.0	595, 590.1, 590.8, 599.0
Pneumonia	485, 486	485, 486
Incisional hernia, disruption of operative wound	551.2, 553.2, 998.3	551.2, 552.2, 553.2, 998.3
Fistula of bladder, postoperative fistula	596.0, 998.6	596.1, 596.2, 619.0
Abdominal pain	785.5	789.0
Other surgical complication	998.9	998.8, 998.9
Cholecystectomy-related Complications		
Cholecystitis, cholangitis	575	575.0, 575.1, 576.1
Cholelithiasis	574	574
Other diseases of gallbladder, biliary ducts	576	575.2 to 575.9, 576.2 to 576.9
Jaundice	785.2	782.4
Gastroenteritis, colitis	009.2	009.0, 009.1
Intestinal obstruction, peritoneal adhesions	560, 568	560, 568
Acute pancreatitis	577.0	577.0
Gastritis, duodenitis	535	535

Two HRSA Programs Benefit the Homeless

• The Health Resources and Services Administration (HRSA) of the US Public Health Service recently announced an award of about $14 million by HRSA to 109 existing health care projects for the homeless in 43 states and the District of Columbia. These 1989 funds will supplement the second year of operation for the projects; $46 million was provided in the first year of operation.

The projects have provided health care to more than 236,000 homeless persons, including many family units with children. Approximately 60 percent of those served are minority group members. In 1989, the number of homeless persons served by the 109 projects is expected to total more than 450,000.

• The "doctors on wheels" project—a major effort to provide outreach health care for the homeless of New York City—is supported by HRSA under its health care for the homeless program. The mobile unit, which is equipped with supplies and a small waiting room, is staffed by health professionals and makes regular stops each week at facilities for the homeless in Manhattan. Sponsored by the New York Children's Health Project, the program is directed by Dr. Irwin Redlener, chief of ambulatory pediatrics at the hospital.

During the three-year fiscal period 1987–89, the project received nearly $1 million from HRSA. The project also operates a second van that serves the homeless in the Brooklyn and Queens sections of New York City. The first mobile unit was acquired with the assistance of singer Paul Simon.

For further information about these projects, contact Frank Sis at HRSA, USPHS, Department of Health and Human Services, 5600 Fishers Lane, Room 14-43, Rockville, MD 20857. Tel: (301) 443-3376 or -3377).

Excessive Burden of End-State Renal Disease among Canadian Indians: A National Survey

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Abstract: Analysis of data from the Canadian National Renal Failure Register indicates that Canadian Natives are at much higher risk for end-stage renal disease (ESRD) than the Canadian population in general. Using two population estimates for the total Native population, the age-standardized incidence rate of newly registered ESRD cases between 1981 and 1986 among Natives was at least 2.5 times (and may be as high as four times) the national rate. Natives were particularly at higher risk for ESRD to diabetes, glomerulonephritis, and pyelonephritis, whereas for the other causes the risk was no different from that of other Canadians. As technologically sophisticated treatment facilities are only available in major urban centers, Native ESRD patients and their families living in remote areas of Canada are faced with major psychosocial disruptions of relocation. (*Am J Public Health* 1989; 79:756–758.)

Introduction

Ethnic differences in the incidence of end-stage renal disease (ESRD) have been observed in some populations. In the United States, Blacks and Hispanics generally have higher rates than Whites.[1–3] Recent reports from the American Southwest also show that the rate among some Indian tribes, particularly the Zuni, exceed those of Whites by several- fold.[4,5] It is not clear if such a phenomenon is also observed in other Amerindian groups across the continent.

This paper reports on the first phase of a multidisciplinary study examining the epidemiological, clinical, and sociocultural dimensions of ESRD among Canadian Indians.

Methods

The Canadian National Renal Failure Register—jointly funded by the Department of National Health and Welfare, Statistics Canada, and the Kidney Foundation of Canada—has collected statistics on chronic renal failure patients since 1981. The register aggregates demographic and clinical data from 75 treatment centers in the nine Canadian provinces with specialized dialysis and transplant programs. It is population-based, voluntary and covers all active renal treatment centers in the nation.[6] Data on all registered patients is updated yearly from medical record reviews in each center to reflect changes in treatment modalities and location, transplant experience and mortality. Patient profiles include age, sex, racial origin, birth date, course of treatment, method of treatment, transplant experience, donor type, cause of graft failure, and cause of death. There is no duplicate counting even as patients are transferred from one center to another. The coding categories and entry criteria are generally compatible with those developed and used internationally by the European Dialysis and Transplant Association.[7]

We requested special computer runs and obtained custom tabulations of all newly registered ESRD cases during the period 1981–86, comparing Native and non-Native patients. In the designation of "racial origin," the National Register distinguishes between Caucasian, Oriental, Black, Native, and other groups. Although the term "racial origin" connotes genetic differences, ascription of individuals to the above categories is more generally made on the basis of "ethnicity," a term which combines genetic (and usually physiognomous) characteristics with socio-cultural, environmental, political, and administrative attributes. The criteria for determining these attributes are subject to considerable interpretation and, at the present time, the National Register does not provide detailed criteria to the individual reporting renal units for establishing "ethnicity" or "racial origin."

To investigate how people of Native ancestry were designated, we interviewed the clinicians responsible for data entry and reviewed the records of one provincial unit. In the central dialysis unit with the highest proportion of Native clients, review of the register and preliminary audits of patient records of all clients (N = 63) revealed that 28 percent were designated as Native and 72 percent non-Native. We searched for documentation of band number (which signifies legal membership in a government recognized Indian group), home community address, indication of need for translation services in a Native language, staff notations indicating self-ascription of Native identity, or staff notes indicating attribution of Native ancestry on the basis of physical characteristics or personal history. None of the patients listed in the registry as non-Native were found to have any information indicating Native ancestry. All but four of the patients categorized as Native by the registry had official Indian band numbers listed; among these four individuals, records and follow-up interviews with the nursing staff indicated that two had described their identity as Metis or non-status Indian and two were identified on the basis of language patterns, home address, and physical appearance as being of Native ancestry.

In Canada, there are several categories of Native Peoples. These include Registered (Status or Treaty) Indians, non-status Indians, Metis, and Inuit (Eskimos). Metis are descendants of mixed Indian-White marriages during the early settlement of the Canadian West. They constitute a distinct cultural and political group in Canada today. According to the 1981 Census of Canada, Registered Indians constituted only 60 percent of all Native Peoples in the country. Since it is not clear as to the true denominator of the Native cases identified by the National Register, two different estimates of the Native population of Canada were used in the

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computation of rates. The Department of Indian and Northern Affairs (DINA) publishes annually the population by age, sex, and residence of all Registered Indians in Canada. The 1984 (representing the mid-point of the study period) Registered Indian population serves as the "A" estimate. This is a low estimate and assumes all Native cases in the register were exclusively Registered Indians. A second, higher estimate of the denominator assumes that the Native ESRD cases belong to all categories of Native Peoples. By dividing the "A" population estimate by 60 percent we obtained an estimate (the "B" estimate) of the total Native population in 1984. It is likely that the true denominator lies somewhere between the "A" and "B" estimates.

As the age structure of the Native population differs substantially from the Canadian national population, the former being much younger, age standardization was performed using the Canadian population as standard. The mid-period population was obtained by averaging the 1981 and 1986 Census of Canada population by age-sex groups. In the comparison of total all-causes incidence rates, the direct method of age-standardization was used. In the comparison of cause-specific rates, standardized incidence ratios (SIR) were computed using the indirect method, due to the much smaller population within each age-sex class. Ninety-five percent confidence intervals were computed for all rates and ratios.[8]

Results

During the period 1981–86, there were a total of 8,432 newly registered cases of ESRD in Canada (5130M, 3302F), of whom 304 were categorized as Native (162M, 142F). Table 1 indicates that at all age-sex groups, the Native rate (23.1/100,000 using the "A" estimate and 13.9/100,000 with the "B" estimate), regardless of whether the "low" or "high" population estimate was used, was higher than the Canadian national rate. Age-standardization resulted in an even higher Native rate, which may be as high as four times and at least 2.5 times the national rate.

The proportion of cases due to diabetes and glomerulonephritis is higher among Natives than all Canadians, whereas the proportion due to renovascular disease and polycystic kidneys is lower (Table 2). In terms of the SIR by cause, Natives are at least at twice the risk as all Canadians for ESRD due to diabetes, glomerulonephritis, and pyelonephritis. Sixty-two percent of Native patients received hemodialysis as the initial treatment compared to 56 percent among all Canadians.

The crude prevalence of ESRD at period end (December 31, 1986) was 22.3/100,000 for all Canadians, and 53.4 ("A" estimate) or 32.0 ("B" estimate) for Natives. The cumulative mortality rate for all newly registered patients during this period was comparable for the two groups: 33.1 percent for all Canadians and 31.9 percent for Natives.

Discussion

This national survey confirms clinical observations and limited mortality and hospital utilization data that a higher burden of ESRD exists among the Native population of Canada. The extent and magnitude of ESRD in the Indian population of Canada has not been adequately assessed. However, several crude estimates based on death certificates and medical care insurance claims suggest that the risk of renal disease is higher among Indians than the general population.

A study of mortality among residents of Indian reserves aged 1–69 in seven provinces from 1977–82 showed that, while kidney diseases were not among the 10 highest ranking causes of death, Indian men had two times and Indian women four times the risk of dying from kidney diseases compared to Canadians nationally.[9] According to claims submitted to the Manitoba Health Services Commission, the provincial health care insurance agency, the hospital separation rate for Manitoba Indians for all genitourinary diseases (ICD-9 codes 580–629) was 15.5/1000 per year, compared to 10.2/1000 for all Manitobans. For the rubrics "nephritis, nephrotic syndrome and nephrosis" (codes 580–589) the Indian rate was 1.4/1000 and the provincial rate was 0.4/1000. The age-standardized ratio for all genitourinary diseases comparing Indian with the provincial total was 2.3.

TABLE 1—Incidence Rate of Newly Registered Chronic Renal Failure: Comparison of Natives with all Canadians, 1981–86

Age-Sex Group	All Canadians	Rate Ratio Native:Canada (A)	95% CI	(B)	95% CI
Male (years)					
0–4	0.71	1.90	0.46, 7.94	1.14	0.28, 4.76
5–14	0.68	3.31	1.45, 7.62	1.99	0.87, 4.57
15–24	2.68	3.46	2.28, 5.29	2.08	1.37, 3.17
25–44	5.65	2.52	1.82, 3.48	1.51	1.09, 2.09
45–64	14.23	4.04	3.15, 5.20	2.43	1.89, 3.12
65+	21.06	3.32	2.31, 4.76	1.99	1.39, 2.85
Subtotal	6.96	2.15	1.84, 2.51	1.29	1.10, 1.51
Female (years)					
0–4	0.38	1.87	0.25, 13.91	1.13	0.15, 8.35
5–14	0.71	4.41	2.11, 9.07	2.65	1.27, 5.44
15–24	2.05	2.76	1.61, 4.71	1.65	0.96, 2.82
25–44	3.35	4.02	2.89, 5.60	2.41	1.74, 3.36
45–64	8.80	6.34	4.87, 8.24	3.80	2.92, 4.95
65+	10.15	5.26	3.51, 7.89	3.16	2.11, 4.73
Subtotal	4.39	3.03	2.56, 3.59	1.82	1.54, 2.15
Total (Crude)	5.66	2.50	2.23, 2.80	1.50	1.34, 1.68

NOTES: All rates are expressed as cases per 100,000 population per year. (A), (B)—See text for definition of different Native Canadian population estimates.

TABLE 2—Etiology of ESRD in Canada, 1981–86: Native vs All Canadians

Cause	All Canadians	Natives	(A) Estimate SIR	(A) Estimate 95% CI	(B) Estimate SIR	(B) Estimate 95% CI
	n = 8432	n = 304				
Glomerulonephritis	24.5%	29.3%	4.1	3.3, 5.0	2.4	1.9, 3.0
Diabetes Mellitus	17.7	25.7	5.3	4.2, 6.6	3.2	2.5, 4.0
Renovascular Disease	13.0	3.3	1.2	0.6, 2.2	0.7	0.3, 1.3
Pyelonephritis	10.6	11.8	3.8	2.6, 5.3	2.3	1.6, 3.2
Polycystic Kidneys	7.2	2.6	1.5	0.7, 3.0	0.9	0.4, 1.8
Drug-Induced Nephropathy	1.8	0.3	0.8	0.0, 4.4	0.5	0.0, 2.8
Others	14.4	14.5	3.3	2.4, 4.4	2.0	1.5, 2.7
Unknown	10.9	12.5	5.3	3.8, 7.2	3.3	2.3, 4.4

NOTE: (A), (B)—See text for definition of different Native Canadian population estimates. SIR = Standardized Incidence Ratios.

The precise causes for this phenomenon are not known but may have been due to the higher prevalence of diabetes mellitus[10,11] known to be present in many Indian groups. Diabetes accounts proportionately for more ESRD cases among Natives than other Canadians. The incidence of impetigo-related post-streptococcal glomerulonephritis is also higher in Native communities, although it has yet to be established that such cases are more likely to proceed to ESRD. It is also interesting that, among the different sites of cancers, Indians are particularly at high risk for kidney cancers.[12,13] It would therefore appear that, for a variety of environmental and genetic reasons, the kidneys of Canadian Natives are highly susceptible to pathologic derangements which may result in ESRD or malignant change.

The lack of uniform and clearly defined criteria for racial origin in the National Register and its confusion with ethnic identity, however, impose some reservation on our conclusions. Many people designated as Native using nongenetic criteria are usually of mixed racial heritage which obviously complicates discussions of genetic predisposition to certain diseases. Furthermore, different renal reporting centers may have greater or lesser sensitivity to sociocultural or environmental criteria for determining Nativeness which may contribute to either over- or under-reporting errors. To assess variability in the designation of ethnicity, current criteria and validity of data on Native ancestry in each of the country's treatment centers need to be documented.

The demonstration of an excess in the burden of ESRD among Canadian Natives has profound implications for health care delivery. In order to survive, patients must depend on technologically sophisticated and professionally controlled treatment provided in urban hospitals. Many Native communities in Canada are located in remote areas, and relocation of Native patients and their family members for medical treatment disrupts social support patterns and creates a host of psychosocial problems.

Further epidemiological analyses will determine ethnic differences in the distribution of risk factors, types, and frequency of complications, survival experience, the impact of urban migration and interaction with medical technology. Concurrently the cultural and socioeconomic basis for discontinuities between service providers' and Native clients' interpretation of treatment alternatives, risk factors, and quality of life will be examined.

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Utilization of Hospital Services by Cardiovascular Patients, Alberta, Canada

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Abstract: Using hospital discharge records, and United States DRG (diagnosis related groups) data, we studied hospital utilization by cardiovascular patients, associated hospital expenditures, and the per capita cost of treating cardiovascular diseases in Alberta, Canada between 1971 and 1986. Expressed in constant 1984 Canadian dollars, the estimated total hospital cost increased from $84 million in 1971 to $131 million in 1986; during this period the Province of Alberta spent about $51 Canadian per resident each year for cardiovascular hospital services. It was noted that rural residents consumed a higher volume of resources per capita than their urban counterparts. A patient origin-destination analysis indicated an increasing dependence of rural patients on urban hospitals for secondary or tertiary care, underscoring the effects of medical technology on referral patterns. (*Am J Public Health* 1989; 79:759–764.)

Introduction

Cardiovascular (CV) disorders are the leading contributor to time spent by Canadians in acute care hospitals. This has been a major concern to hospital planners because of the growing emphasis on health care cost containment. Both the total and per capita costs of CV disorders are difficult to estimate because Canadian hospitals are funded under global budget arrangements, and individual patient accounts are not generally maintained.

In most Western industrialized countries, the advent of a plethora of sophisticated techniques and instruments for treating CV disorders over the past two decades has changed the nature of their hospital care. These changes have been accompanied by declines in CV disease incidence rates. Although advanced technology has been associated with higher health care costs, the impact of technological changes cannot be assessed without proper cost estimates. Moreover, the health care cost saving due to the recent emphasis on life-style modification cannot be substantiated without reliable estimates of the total costs involved.

Canadian mortality rates attributed to ischemic heart disease increased until the 1960s and then declined approximately 1.7 per cent for males and 2.4 per cent for females annually during the 1970s;[1] age-adjusted mortality rates for ischemic heart disease declined by 16.4 per cent between 1969 and 1977,[2] as compared to a 20.7 per cent decrease in the US.[3] It is generally conceded that this decline is real rather than an artifact of diagnostic variations, increased awareness, or changes in disease classification. Specific causes of declining mortality rates have been predominately related to changes in CV risk factors and, secondarily, to improvements in and accessibility of coronary medical services.[1,3]

In Canada, approximately 8.5 per cent of the gross national product is devoted to health care.[4] Capital and operating hospital expenditures are normally provided by the provincial governments and individual patients are not billed directly. Physicians and surgeons who practice in hospitals bill the provincial health care insurance plans directly on a fee-for-service basis, virtually eliminating financial barriers to health services. Removal of financial deterrents to seeking health services, however, has resulted in increased hospital and medical care utilization and associated expenditures.

Alberta, one of 10 Canadian provinces, has a population of 2.4 million. Serving this population are 123 general acute care hospitals providing an average of 5.5 beds per 1,000 persons. On an average annual basis, Albertans are admitted to hospital 0.19 times and spend 1.4 days in hospital. The total health care per capita cost in 1984 was about C$1,721, and the government allocated approximately C$698 per capita for Alberta hospitals.[5] During the study period (1971–86), all hospitals in Alberta subscribed to the Professional Activity Study (PAS) hospital discharge abstract system, generating about 400,000 separation records per year. The files are fully computerized and available for research and planning purposes.

Using hospital discharge data, Newman[6] reported that CV disease ranked as the fourth and sixth most important cause of hospitalization for Alberta males and females, respectively, in 1980, but was the outstanding cause of time spent in Alberta acute care hospitals. This high volume of acute care hospitalization has rendered CV disease a major health care expenditure.

The purpose of this study was to explore utilization and resource consumption by CV patients in acute care hospitals in the Canadian province of Alberta. Specific objectives were:

- An examination of global patterns and trends in acute care hospital utilization from 1971 and 1986;
- An historical comparison of total, per capita, and per discharge hospital costs;
- A comparison of utilization rates over time among geographic regions of Alberta; and
- An exploration of the care-seeking patterns of patients.

Methods

Data Source

The primary source of the clinical data was PAS records, a patient discharge abstract system developed for acute care hospitals by the Commission on Professional and Hospital Activities, Ann Arbor, Michigan. Variables coded on the discharge abstract include age, sex, date of admission and discharge, length of stay, primary, secondary and tertiary diagnoses, operation codes, discharge status, patient's home address and other pertinent variables. The current disease coding scheme used for diagnoses and operation codes is the

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ICD-9-CM; diagnosis related group (DRG) codes are available only for separations generated in 1984 and thereafter.

The CV disease diagnoses were organized according to the Major Disease Category of the revised ICD-9-CM version of the DRG classification system.[7] Abstract files were extracted for all Alberta residents discharged from acute care hospitals with CV disorders between 1971 and 1986; DRG codes were assigned retrospectively for the discharges generated before 1984. During the course of this retrospective assignment of DRG codes for pre-1984 discharges, it became apparent that differences in coding schemes precluded the achievement of perfectly comparable DRGs. After the revised ICD-9-CM circulatory DRGs were finally adjusted, we constructed similar DRG groups for the years 1971 to 1978 inclusive, incorporating the H-ICDA coding scheme for 1971 to 1973 and the H-ICDA-2 coding scheme for 1974 to 1978. Initial cross-tabulations and comparisons with the Canadian Diagnostic Listing system indicated that the resulting DRG groups were relatively similar and representative of overall CV disorders (The extensive Coding Guides used for this project are available from the authors on request). Since this study was designed to study general patterns and trends of CV utilization and did not involve epidemiological assessment of individual disease categories, we believe that this limitation does not imply any significant compromise in the findings and conclusions.

We examined the use of hospital resources incorporating as proxies the DRG weight values used by the US government, henceforth referred to as the weighted DRG (See Appendix I). One unit of the DRG weight is equivalent to the resources required by an average American hospitalization episode. The weighted sum of discharges provided a measure of relative resource consumption by a particular group of hospital patients. The US Prospective Payment Assessment Commission[8] provided the DRG weights in effect during the fiscal year 1987 and average charges and geometric mean lengths of stay based on 1984 data.

The other major data source was the census data released by Statistics Canada for the years 1971, 1976, 1981, 1986, and estimates of the Alberta population for the intercensus years. The Alberta government's report on hospital expenditures for the year 1984 was also used.

Analyses

Total number of discharges, total number of patient-days, and total weighted DRGs were compiled for each year from 1971 to 1986 for medical, surgical, and combined CV DRGs. By using the average hospital cost per unit of the weighted DRG, experienced in 1984 by all Alberta hospitals for all discharges (CV and non-CV), a particular group's weight values were converted into CV hospital service cost estimates in 1984 constant Canadian dollars. By dividing these basic measures by the corresponding population figures, population-based hospital utilization rates were derived including the number of discharges per 1000 person-years, the number of patient-days per 1000 person-years, and the cost per person-year in 1984 Canadian dollars. By dividing the total patient-days, weighted DRGs and costs by the corresponding total number of discharges, per-case measures such as average length of stay and average weighted DRG per discharge were also derived. Alberta was divided into metropolitan, suburban, and rural areas and annual per capita and per case utilization rates were computed and compared for the three regions.

The final step of the analysis centered on an investigation of the care-seeking behaviors of patients with CV disorders. Instead of using the traditional discharge and patient-day indicators as measures of resource use, the weighted DRG was used for a more direct assessment of resource consumption and allocation patterns. Relevance indices[9,10] measure the extent to which CV patients of a geographic area were dependent on the resources of a hospital group; commitment indices[9,10] measure the extent to which selected hospital(s) allocated CV resources to patients originating from a particular geographic area (see Appendix II for definitions of these indices). The indices were obtained for each study year and historical patterns were identified.

Results

Hospital Utilization: All DRGs

If the average sum per DRG weight unit in 1984 paid by the US government (US$4,155) is valid for Alberta cases, this value may be applied to the Alberta weighted DRG. As total Alberta weighted DRGs increased from 25,300 in 1971 to 39,400 in 1986, the corresponding cost estimates would be US$105 million and US$164 million, respectively, for 1971 and 1986, and the total cost estimate over the years would be US$2.07 billion. These amounts would have been paid to hospitals for CV services if the Alberta hospital system had adopted the US DRG prospective payment system during the study period. However, the validity of this approach can be seriously challenged as Alberta hospitals operate very differently from their US counterparts and the data of this study included the whole population while the US data were based primarily on those aged 65 or over (all Medicare and a small portion of Medicaid discharges). Thus, as noted previously, a somewhat different approach was adopted to address the above concerns.

Cost per DRG Weight Unit

During 1984, Alberta acute care hospitals generated 439.6 thousand discharges yielding a total of 337.5 thousand weighted DRG units. Thus the average weighted DRG value per discharge for all Alberta discharges (both CV and non-CV cases) was 0.77 units, much lower than the theoretical value of 1.0. In order to compare the average weighted DRG between the US and Alberta, selection criteria were applied to the two data sets. For US data, only those DRG codes which were applicable to persons over age 65 were selected; for Alberta data, patients under age 65 were excluded, thus making the two data sets composed of groups 65 years and over. The average weight values per discharge for the elderly were 1.11 and 1.03 for the US and Alberta, respectively. The lower Alberta figure suggests that Alberta hospitals experienced relatively less resource-intensive case-mixes in 1984. It is interesting to note that DRG127 (Heart failure and shock) ranked the highest of total weighted DRGs for the elderly group in both US and Alberta; this particular DRG accounted for 4.3 per cent and 3.9 per cent of the total elderly weighted DRGs in 1984 (Alberta data are available from the authors on request). During 1984, Alberta hospitals spent about C$1,121 million to operate 123 acute care hospitals resulting in C$3,321 per unit of weighted DRG.

Hospital Utilization: CV DRGs

Historical trends associated with hospital utilization by CV DRGs (DRG Codes 103–145, see Appendix I for listing) are presented in Table 1 for per capita rates and in Table 2 for per case rates. During the 16-year period, the Alberta

TABLE 1—Per Capita Utilization Rates for Medical and Surgical CV DRGs, Alberta Hospitals, 1971–86

	Medical DRG			Surgical DRG			All CV DRG		
Year	Dis*	Pt-Days#	Cost†	Dis	Pt-Days	Cost	Dis	Pt-Days	Cost
71	11.5	158	34	2.6	37	17	14.1	195	52
72	12.3	167	36	2.6	40	18	15.1	207	55
73	11.0	139	33	2.6	39	18	13.7	177	52
74	11.1	134	33	2.5	36	18	13.5	169	52
75	10.8	127	32	2.5	35	19	13.3	162	51
76	10.8	122	33	2.5	36	20	13.3	158	52
77	10.4	112	32	2.4	32	19	12.8	144	51
78	10.0	107	31	2.4	36	21	12.4	142	51
79	9.7	103	30	2.3	34	20	12.0	137	49
80	9.1	92	28	2.2	32	19	11.3	124	47
81	9.0	89	28	2.1	30	18	11.0	119	46
82	9.3	89	28	2.2	31	19	11.4	120	47
83	9.6	89	29	2.2	31	20	11.8	121	49
84	9.9	92	31	2.4	33	22	12.3	124	52
85	10.3	92	32	2.6	34	25	12.9	126	57
86	10.1	90	31	2.5	33	24	12.6	123	55
71–86	10.2	109	31	2.4	34	20	12.6	143	51

*Discharge per 1000 person-year
Patient-days per 1000 person-year
†Cost per person-year in 1984 Canadian $

population increased by 44.3 per cent, but the number of discharges increased by only 29.6 per cent, indicating an overall decline of hospital admission rates and/or a decline of CV incidence. The per capita discharge rate declined 10.6 per cent from 14.1 per 1,000 person-years in 1971 to 12.6 in 1986. The total number of patient-days declined 8.6 per cent and the per capita rate decreased as much as 36.0 per cent. As these measures do not take into account case-mix variation, these statistics do not reflect necessarily the resource consumption by CV patients. Although it is not perfect, the resource consumption estimate based on DRG classification provided a more accurate picture.

Using the average cost per weighted DRG unit in 1984 (C$3,321), the total estimated cost increased from C$84 million in 1971 to C$131 million in 1986, a gain of 56.0 per cent which is slightly greater than the population growth (44.3 per cent); the per capita cost for CV hospital services increased minimally from C$52 to C$55. Thus it appears that decreased CV morbidity, supported by declines of discharge and mortality rates, did not result in a significant reduction of the per capita cost; per capita cost was more highly associated with health care system variables than with epidemiological variables. There had been a gradual decline in the per capita cost for the years 1976–81, corresponding to the time period when Alberta experienced economic growth. With the recession and stabilization of the population since 1982, the per capita rates gradually increased to the 1972 level where the highest rates were observed, suggesting the effects of health service supply on utilization.

TABLE 2—Per Case Utilization Rates for Medical and Surgical CV DRGs, Alberta Hospitals, 1971–86

	Medical DRG		Surgical DRG		All CV DRG	
Year	Days*	DRG#	Days	DRG	Days	DRG
71	13.7	0.90	14.2	1.97	13.8	1.10
72	13.6	0.89	14.3	1.99	13.7	1.09
73	12.5	0.91	14.7	2.10	13.0	1.14
74	12.1	0.91	14.6	2.23	12.5	1.15
75	11.8	0.91	14.2	2.28	12.2	1.16
76	11.4	0.91	14.1	2.37	11.9	1.19
77	10.7	0.91	13.4	2.41	11.2	1.19
78	10.7	0.92	14.7	2.56	11.4	1.24
79	10.7	0.93	14.5	2.51	11.4	1.24
80	10.2	0.93	14.5	2.63	11.0	1.27
81	9.9	0.93	14.6	2.66	10.8	1.25
82	9.6	0.92	14.4	2.67	10.5	1.25
83	9.3	0.92	14.1	2.73	10.2	1.26
84	9.3	0.93	13.7	2.77	10.1	1.28
85	8.9	0.93	12.9	2.91	9.7	1.33
86	8.9	0.94	13.3	2.89	9.7	1.32
71–86	10.7	0.92	14.1	2.51	11.4	1.22

*Patient-days per discharge
#Weighted DRG per discharge

For per case rates (Table 2), the average length of stay declined 29.7 per cent, from 13.8 days per discharge in 1971 to 9.7 days in 1986; however, the average weighted DRG values gradually increased from 1.10 to 1.32 (+20.0 per cent), suggesting a substantial shift of case mix toward more resource-intensive admissions over the years. Thus, the overall decline of per capita discharge and patient-day rates coincided with a shorter average length of stay and increased case mix complexity; however, these factors had a mutually canceling effect resulting in a relatively stable per capita cost.

Medical DRGs

For medical DRGs (DRG Codes 121–145, see Appendix I for listing), the discharge rate declined 12.2 per cent while the patient-day rates experienced a remarkable decline of 43.0 per cent. Overall, the per capita cost for medical discharges declined about 8.8 per cent. The decline in the patient-day rate suggests an overall productivity gain in medical DRGs. This observation is also supported by the per case rates as shown in Table 2. During the 16-year period, the case mix complexity, as measured by average weighted DRG per discharge, increased slightly (4.4 per cent) while the average length of stay declined 35.0 per cent, suggesting that Alberta hospitals were admitting patients with greater case mix complexity but keeping them a shorter length of time on average.

Surgical DRGs

For surgical DRGs (DRG codes 103–120, see Appendix I for listing), per capita discharge rates declined from 2.6 per 1,000 person-years in 1971 to 2.1 in 1981 but increased to 2.5 in 1986 (Table 1). The patient-day rates exhibited a similar pattern but declined slightly (−10.8 per cent). In sharp contrast to these rates and the medical DRG per capita cost, the overall per capita surgical cost increased 41.2 per cent, reflecting the more resource-intensive case mixes in later years. During the 16-year period, surgical CV cases gradually became more resource-intensive suggesting the effects of a general shift to higher technology and greater sophistication of surgical methods. The portion of all CV per capita cost attributable to surgical DRGs increased from 32.7 per cent in 1971 to 43.6 per cent in 1986, suggesting a substitution effect of medical DRGs by the surgical DRGs. The case mix complexity of surgical DRGs increased dramatically during the period: the average weighted DRG value increased from 1.97 to 2.89, an increase of 46.7 per cent. Alberta hospitals, however, were able to reduce the average hospital stay about one day (from 14.2 to 13.3), indicating the effects of medical technology and tighter bed supply.

Geographic Factors

The results of a regional comparison are summarized in Table 3 for all CV DRGs. Rural discharge rates for all CV DRGs were about twice those of the metropolitan areas,

TABLE 3—Per Capita Utilization Rates for All CV DRGs by Origin, Alberta Hospitals, 1971–86

Year	Discharges*			Patient-Days#			Cost†		
	Metro-politan	Sub-urban	Rural	Metro-politan	Sub-urban	Rural	Metro-politan	Sub-urban	Rural
71	10.6	12.5	19.9	158	181	255	42	46	67
72	11.0	14.1	21.5	156	204	285	44	49	73
73	10.1	12.9	19.4	139	171	238	42	47	68
74	9.9	12.7	19.3	131	168	229	41	49	68
75	9.7	12.0	19.1	128	150	219	42	46	67
76	9.7	12.9	18.9	123	156	212	43	50	67
77	9.3	12.6	18.3	110	146	196	42	50	66
78	9.1	12.0	17.7	113	135	191	43	48	66
79	8.7	11.8	17.2	110	135	180	40	48	65
80	8.3	10.9	16.1	101	120	162	39	45	61
81	8.1	10.2	15.8	95	112	158	38	41	60
82	8.6	11.1	16.0	100	115	153	40	44	60
83	8.9	11.3	16.5	101	120	153	41	47	63
84	8.8	12.0	18.0	98	119	170	42	49	71
85	9.2	12.1	19.3	98	119	174	45	52	78
86	9.0	11.8	18.8	98	117	166	45	50	75
71–86	9.2	11.9	18.1	113	137	192	42	47	67

*Number of discharges per 1000 person-year
#Number of patient-days per 1000 person-year
†Cost per person-year in 1984 Canadian $

TABLE 4—Per Capita Utilization Rates for Surgical DRGs by Origin of Patient, Alberta Hospitals, 1971–86

Year	Discharges*			Patient-Days#			Cost†		
	Metro-politan	Sub-urban	Rural	Metro-politan	Sub-urban	Rural	Metro-politan	Sub-urban	Rural
71	2.7	2.7	2.5	39	36	35	18	16	16
72	2.8	2.7	2.8	40	34	42	19	15	18
73	2.7	2.5	2.6	40	32	39	19	16	18
74	2.4	2.4	2.5	36	32	37	18	17	18
75	2.5	2.3	2.5	36	27	36	19	17	19
76	2.5	2.5	2.5	37	28	35	21	18	19
77	2.3	2.6	2.5	33	29	33	20	19	19
78	2.5	2.3	2.4	37	29	35	22	18	20
79	2.3	2.4	2.4	34	33	34	19	19	20
80	2.1	2.4	2.2	32	32	31	19	19	19
81	2.1	2.1	2.1	31	27	29	19	16	18
82	2.2	2.1	2.2	33	25	30	20	17	19
83	2.2	2.1	2.3	32	30	31	20	18	20
84	2.3	2.2	2.7	32	27	37	21	19	25
85	2.4	2.5	2.9	32	30	37	24	22	28
86	2.4	2.2	2.7	33	27	36	23	20	27
71–86	2.4	2.3	2.5	34	29	35	20	18	20

*Number of discharges per 1000 person-year
#Number of patient-days per 1000 person-year
†Cost per person-year in 1984 Canadian $

whereas suburban rates were slightly higher than metropolitan rates. Rural discharge rates changed very little over the years while metropolitan residents experienced a slight decline. Throughout the study period, rural residents used hospital beds 52-83 per cent more than their metropolitan counterparts on a per capita basis, resulting in higher per capita CV costs for rural residents. Thus it is clear that rural residents of Alberta were consuming a much higher volume of hospital resources over the study years. This is consistent with findings of other utilization studies conducted in Alberta;[11,12] obviously these differences cannot be explained by the medical need model. It appears that the most probable factors accounting for these differences are associated with system-related determinants of utilization such as multiple admissions for the same disease episode for rural residents through referral networks; higher bed availability in rural communities;[10] more cautious practice by the physicians for patients originating from rural areas; and lack of alternative services in rural areas. To substantiate this interpretation, utilization rates for surgical DRGs were tabulated separately (see Table 4). As expected, there were very few discernible systematic variations in per capita utilization rates among the three geographic areas in Alberta. Thus it may be concluded that the rural-urban variation is mostly due to system-related determinants of health service utilization.

Patient Flow Patterns

The results of the patient origin-destination analysis are summarized in Table 5. The indices were derived using total weighted DRGs attributable to each segment of patient origin and destination combination, partially controlling the variation in resource needs due to case mix differences. From a community perspective, results indicated a markedly increasing tendency for those in rural areas in Alberta to depend on metropolitan hospital resources. For example, the relevance index of metropolitan hospitals to the rural area, RI(M,R), or the dependence of rural residents on metropolitan hospitals, increased from 26.3 per cent in 1971 to 46.8 per cent in 1986; almost one-half of the weighted DRGs generated by rural residents represented resources consumed in metropolitan hospitals in 1986. Conversely, the relevance index of rural hospitals to rural areas, RI(R,R), declined from 71.8 per cent to 48.1 per cent. Some of this shift may be attributed to technological changes as highly sophisticated facilities and interventions by specialists are available only in urban large hospitals. The relevance index of metropolitan hospitals to suburban area, RI(M,S), also increased from 19.9% to 31.5%. Thus it is obvious that both rural and suburban areas of Alberta are increasingly depending on metropolitan hospital CV services.

From a provider perspective, commitment indices revealed an inverse relation between the level of care available within a hospital group and the commitment of that group to

TABLE 5—Relevance and Commitment Indices Based on Weighted DRGs, Alberta Hospitals, 1971–86

Year	RI(M,R)*	RI(R,R)	RI(M,S)	CI(M,M)	CI(M,R)	CI(S,R)
71	26.3	71.8	19.9	75.7	21.3	12.6
72	26.9	70.9	18.4	74.6	22.5	13.4
73	29.4	68.3	21.6	73.0	23.7	13.7
74	30.3	67.5	26.3	71.7	24.0	14.0
75	33.0	64.6	29.1	70.4	25.2	15.9
76	33.9	63.7	27.1	70.6	25.0	14.0
77	34.7	62.7	29.6	69.5	25.5	15.0
78	37.0	59.7	29.5	69.1	26.3	18.3
79	37.9	58.9	27.9	67.7	27.6	16.9
80	38.4	58.1	29.1	68.2	27.0	18.2
81	39.0	57.6	27.0	68.1	27.6	18.3
82	40.1	55.7	30.3	68.3	26.8	21.2
83	41.5	54.6	31.1	67.5	27.4	19.2
84	45.0	49.7	31.8	63.6	31.3	25.9
85	46.6	48.9	38.1	62.5	31.6	25.0
86	46.8	48.1	35.1	63.3	31.4	26.7
71–86	37.4	59.1	29.1	68.3	27.0	18.9

*Character within bracket denotes hospital group and geographic area using the following:
M (Metro), S (Suburban), R (Rural).
RI(M,R) Relevance Index of Metro Hospitals to Rural Areas
RI(R,R) Relevance Index of Rural Hospitals to Rural Areas
RI(M,S) Relevance Index of Metro Hospitals to Suburban Areas
CI(M,M) Commitment Index of Metro Hospitals to Metro Areas
CI(M,R) Commitment Index of Metro Hospitals to Rural Areas
CI(S,R) Commitment Index of Suburban Hospitals to Rural Areas

patients within their respective communities. Between 1971 and 1986, metropolitan and suburban hospitals devoted decreasing amounts of CV resources to their respective residents and, concurrently, significantly increased their allocation of CV resources to rural patients. Metropolitan hospitals allocated 75.7 per cent of their resources to metropolitan residents in 1971, but the proportion decreased to 63.3 per cent in 1986. The commitment index of metropolitan hospitals to rural hospitals, CI(M,R), increased from 21.3 per cent in 1971 to 31.4 per cent in 1986. Similar trends were observed for suburban hospitals.

Selected DRGs

Cost estimates for relatively common medical and surgical DRGs are presented in Table 6 suggesting disease specific trends. As fixed costs are assigned to each of the DRG categories, the trends are identical with the number of admissions in each category. For example, DRG122 (Circulatory Disorders with Acute Myocardial Infarction without complications, alive) and DRG127 (Heart failure and shock) increased 56.2 per cent and 48.9 per cent, respectively, slightly more than the rate of population increase (44.3 per cent) whereas DRG130 (Peripheral vascular disorders, 70+ and/or Complication or Comorbidity) increased only 31.6 per cent. The number of coronary bypass procedures (DRG106–7) in 1986 was 5.6 times that of the 1971 value, while valve procedures (DRG104–5) more than doubled, indicating the major causes of the surgical DRG cost increase. DRG108–12 (Cardiothoracic and vascular procedures) almost tripled, while DRG119 (Vein ligation and stripping) declined 41.4 per cent. For the 16-year period, the largest portion of total CV costs was attributed to DRG108–12 (17.8 per cent), DRG127 (10.8 per cent), and DRG122 (9.0 per cent). It is estimated that valve and bypass procedures combined cost Alberta taxpayers in 1986 about C$15 million for hospital services alone. It should be recognized that these figures do not include other costs associated with CV disorders such as physician and out-of-hospital services.

TABLE 6—Total Cost Estimates for Selected DRGs, Alberta Hospitals, 1971–86 (Unit = Million in 1984 Canadian $)

	DRG CODES*								
Year	DRG-122	DRG-127	DRG-130	DRG-132	DRG-134	DRG-104-5	DRG-106-7	DRG-108-12	DRG-119
71	7.3	9.0	3.8	9.8	5.3	2.3	1.7	10.3	5.8
72	7.5	9.8	3.9	11.1	6.4	2.2	2.4	11.1	6.0
73	8.1	9.9	4.1	7.3	4.5	2.1	3.0	12.6	4.9
74	8.6	9.9	4.3	7.4	4.6	2.7	2.8	12.9	4.5
75	8.3	10.0	4.4	6.9	4.7	3.3	3.2	14.1	4.5
76	8.7	10.8	3.9	6.4	4.6	3.0	4.7	15.9	4.1
77	8.6	11.2	4.1	6.2	4.1	2.7	5.1	16.7	3.8
78	8.5	11.2	4.2	5.1	4.0	3.6	5.9	19.2	3.3
79	10.0	11.5	4.2	5.3	3.6	3.1	7.3	17.0	3.5
80	9.6	11.1	3.9	4.8	3.2	4.4	7.1	17.6	3.1
81	9.7	11.5	3.9	4.8	3.3	4.0	8.0	16.6	2.9
82	9.9	11.3	4.0	5.2	3.7	4.3	7.3	20.6	3.4
83	10.2	11.7	4.4	4.7	3.8	4.9	7.8	22.4	3.3
84	11.2	12.7	4.5	4.3	3.6	4.1	7.2	26.8	3.3
85	11.5	13.3	5.4	4.3	3.9	4.7	9.9	31.7	3.3
86	11.4	13.4	5.0	4.0	3.8	5.3	9.5	28.7	3.4
71–86	149.0	178.3	68.0	97.6	67.1	56.8	92.9	294.0	63.2

DRG122 Circulatory disorders with acute myocardial infarction, without complication, alive
DRG127 Heart failure and shock
DRG130 Peripheral vascular disorders, 70+ and/or complication or comorbidity (CC)
DRG132 Atherosclerosis, 70+ and/or CC
DRG134 Hypertension
DRG104-5 Cardiac valve procedures
DRG106-7 Coronary bypass procedures
DRG108-12 Cardiothoracic and vascular procedures
DRG119 Vein ligation and stripping

On the basis of these findings, we draw the following conclusions:

- The overall downward trends in per capita CV utilization in Alberta were consistent with the overall decreasing Canadian hospital utilization rates and decreasing CV morbidity and mortality in developed countries; however, this decline did not contribute to a decrease in the per capita hospital cost which remained relatively stable at about C$51. The introduction or increased use of more resource intensive surgical technology appears to have offset cost savings due to declining utilization rates;
- High per capita hospital utilization rates were primarily a rural phenomenon for medical DRGs, reflecting the effects of the availability of more beds in rural areas, technological changes, referrals, more cautious medical practice patterns by physicians for rural patients, and perhaps the effect of tighter bed control implemented by the provincial government during the latter part of the study period; however, no discernible variation in utilization rates was found for surgical DRGs;
- A substitution effect of the medical DRGs by the surgical DRGs occurred during the study period which significantly influenced the utilization of CV resources in view of the high costs associated with the major surgical DRGs;
- The increasing dependency of non-metropolitan patients on metropolitan hospitals' CV resources, combined with the increased use of metropolitan hospital resources by the elderly identified in other studies,[11] could partially account for the perceived bed shortage and long waiting lists for elective surgery experienced by metropolitan residents during the latter part of the study period. This would suggest a need for shifting resources to urban hospitals;
- This analysis provides some baseline information with which to measure the impact of health promotion programs especially if Canadian provinces adopt a DRG-based reimbursement scheme for financing and managing hospital services.

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APPENDIX I

Cardiovascular DRG Weights*

CODE	Description	Weight
103	Heart Transplant	#
104	C Valve Proc with Pump & Cath	7.3151
105	C Valve Proc with Pump & without Cath	6.3388
106	Coronary Bypass with Cath	5.3324
107	Coronary Bypass without Cath	4.6608
108	Cardiothoracic Proc, except Bypass, with Pump	4.7803
109	Cardiothoracic Proc without Pump	4.3579
110	Major Reconst Vascular Proc, 70+ and/or CC	3.3118
111	Major Reconst Vascular Proc, 69− without CC	2.4549
112	V Proc Except Major Reconst Proc	2.2055
113	Amputation except Upper Limb and Toe	2.5345
114	Upper Limb & Toe Amputation	1.8918
115	Permanent C Pacemaker Implant with AMI or CHF	4.1727
116	Permanent C Pacemaker Implant without AMI or CHF	2.9868
117	C Pacemaker Repl & Rev Except Pulse Gen Repl onl	1.2989
118	C Pacemaker Pulse Gen Replacement only	1.9224
119	Vein Ligation and Stripping	0.9164
120	Other Circ Sys Operation Room Procedure	2.2577
121	Circ Dis with AMI & CV Comp, Alive	1.7687
122	Circ Dis with AMI without CV Comp, Alive	1.3267
123	Cir Dis with AMI, Expired	1.3522
124	Circ Dis Except AMI, with Cath & Complex Diag	1.2551
125	Circ Dis except AMI, with Cath, without Complex	0.7265
126	Acute & Subacute Endocarditis	2.9836
127	Heart Failure	1.0098
128	Deep Vein Thrombophlebitis	0.8456
129	Cardiac Arrest	1.7199
130	Peripheral Vascular Dis, 70+ and/or CC	0.8251
131	Peripheral Vascular Dis, 69− without CC	0.6705
132	Atherosclerosis, 70+ and/or CC	0.8037
133	Atherosclerosis, 69− without CC	0.7049
134	Hypertension	0.6363
135	C Congenital & Valvular Dis, 70+ and/or CC	0.8937
136	C Congenital & Valvular Dis, 18–69, without CC	0.7525
137	C Congenital & Valvular Dis, 0–17	0.6315
138	C Arrhythmia & Conduction Dis, 70+ and/or CC	0.8136
139	C Arrhythmia & Conduction Dis, 69− without CC	0.6514
140	Angina Pectoris	0.6894
141	Syncope & Collapse, 70+ and/or CC	0.6187
142	Syncope & Collapse, 69− without CC	0.5335
143	Chest Pain	0.5893
144	Other Circ Diag with CC	0.1160
145	Other Circ Diag without CC	0.8475

List of Abbreviations Used

18–69	Ages 18 through 69 years, inclusive
69−	Ages 69 years or less
70+	Ages 70 years or over
CODE	Description
AMI	Acute myocardial infarction
C	Cardiac
Cath	Cardiac catheterization
CV Comp	Cardiovascular complication
CC	Complication or comorbidity
CHF	Congestive heart failure
Circ	Circulatory
Diag	Diagnosis
Dis	Disorders
Proc	Operation room procedures
Reconst	Reconstructive
Repl	Replacement

*FY'87 DRG weights, provided by US Prospective Payment Assessment Commission.

No weight was established for Heart transplant (DRG103).

APPENDIX II

The Definitions of Relevance and Commitment Indices

The following formulae were used for the computation of relevance (RI) and commitment (CI) indices. This is a modification of common formulae found in the literature:[9,10] the number of admissions or patient days used as a utilization measure was replaced by the weighted DRG values.

$$RI(H,A) = WDRG(H,A)/WDRG(A)$$

$$CI(H,A) = WDRG(H,A)/WDRG(H)$$

where H denotes one or a group of hospital(s);
A a geographical area;
WDRG(H,A) total weighted DRG values generated at a hospital(s) H by the residents of a geographic area A;
WDRG(A) total weighted DRG values generated at all hospitals by the residents of an area A;
WDRG(H) total weighted DRG values generated at a hospital H by the residents of all areas.

The relevance index may be interpreted as the dependence of an area A on a particular hospital H for health services in terms of resource needs, or the market share of area A by the hospital H. The commitment index may be interpreted as a commitment or allocation of resources by a hospital H to serve an area A.

The Association of Age-Related Macular Degeneration and Lens Opacities in the Aged

INGRID Y. LIU, MD, MPH, LON WHITE, MD, MPH, AND ANDREA Z. LACROIX, PHD

Abstract: Data from 3,087 persons age 45 or older in the National Health and Nutrition Survey, 1971–74, showed that subjects with lens opacifying disease had an increased odds for age-related macular degeneraton (AMD) compared to those who had no lens opacities. The crude odds ratio for aphakic patients was 4.6 (95% CI = 2.5, 8.6). The association remained after controlling for age, sex, and systolic blood pressure (a common risk factor) in a logistic regression model. These data are consistent with the hypothesis that light-induced damage may contribute to both lens and retinal disease and suggest that cataract extraction without implantation of ultra-violet/blue light absorbing intraocular lens may place subjects at increased risk of AMD. (*Am J Public Health* 1989; 79:765–769.)

Introduction

Age-related macular degeneration (AMD) and cataracts are two of the most common causes of visual impairment in older persons. Differing sampling methods and diagnostic criteria make prevalence estimates difficult, but the World Health Organization estimated in 1982 that 17 million persons in the world were blind as a result of cataracts. Ferris notes that AMD was the leading cause of registered blindness in the United States and Great Britain as early as 20 years ago.[1] While each disease affects a separate structural component of the eye, their dramatic increase in prevalence with aging suggests common or related processes may be responsible for both conditions.

Although the pathogenesis of neither condition is well understood, mechanisms that involve light-mediated damage have been theorized for both.[2–4] According to these theories, damage to the lens and/or retina occurs when the body's defenses, such as antioxidant stores and enzyme systems, fail or are overwhelmed. Light has been shown to induce retinal lesions in many mammals.[5–7] A role for free-radical-mediated damage in the development of cataracts is supported by the finding of deficient superoxide dismutase activity in cataractous human lenses.[8]

Since light must pass through the lens to reach the retina, cataracts may actually retard the development of AMD. The yellow pigmentation that characterizes nuclear cataracts may filter out blue light that is felt to be especially damaging.[9] Zigman and Collier noted in the diurnal gray squirrel that the yellow pigmented lens filtered near-ultraviolet (UV) radiant energy and protected the retina from physical and functional damage.[5]

There have been very few population-based studies examining this theory. Sperduto, *et al*, found no association between cataracts and AMD in the Framingham Study when the various age-related lens changes were pooled.[10] On reexamination of the same data, when specific types of cataracts were studied, they found a positive association between AMD and cortical changes and a negative association with nuclear sclerosis.[11]

An interesting and important corollary, if some lens opacification is protective of the macula, is that an increased risk may be imposed by surgical removal of cataractous lens. The Report of the Cataract Panel in Vision Research—A National Plan: 1983–87 suggests that aphakic eyes may be at special risk for retinopathy and maculopathy.[12]

From the Epidemiology, Demography, and Biometry Program, National Institute on Aging. Address reprint requests to Ingrid Y. Liu, MD, MPH, EDBP, National Institute on Aging, Room 618, Federal Building, 7550 Wisconsin Avenue, Bethesda, MD 20892. This paper, submitted to the Journal August 5, 1988, was revised and accepted for publication December 28, 1988.

The current study uses data from the first National Health and Nutrition Examination Survey (NHANES-I) to examine whether there is an association between the occurrence of AMD and lenticular opacities and whether aphakics compose a special risk group for AMD.

Methods

NHANES-I was a cross-sectional study of a probability sample (and partial oversample) of the civilian noninstitutionalized coterminus US population, designed to assess the prevalence of certain conditions. Details of the design and conduct of the survey are available elsewhere.[13] The visual component included a short ocular history and detailed examination for visual acuity, motility, and pathology, including tonometry and dilatation. Right and left eyes were examined and rated separately.

For this study, subjects having "senile macular degeneration," "senile disciform macular degeneration," or "senile circinate macular degeneration" were classified as having AMD, while all other persons served as controls (Table 1). A diagnosis of macular degeneration required pathological findings of "loss of macular reflex, pigment dispersion and clumping, and drusen"[14] as well as visual acuity of 20/25 or worse (best corrected) that was felt by the examiner to be due to the disease.

Lens conditions were grouped into four mutually exclusive groups—aphakic eyes, eyes with cataract and associated visual impairment, eyes with lens opacities but no visual impairment, and eyes with no lens opacifying disease (reference group). The cataract group was formed from all immature, intumescent, mature, hypermature, and morgagnian cataracts which the examiner felt were responsible for vision loss to the level of 20/25 or more (best corrected). The definitions for macular degeneration, cataracts, and aphakia are all consistent with those used by Klein in previous analyses of this data.[15]

The study population was restricted to subjects 45 years of age or older. The odds of having AMD for each of the lens opacifying conditions was calculated for right eyes and left eyes independently, as well as for persons with the condition in either eye. The reference group for the latter "person-measure" analysis was defined as those individuals with normal lenses bilaterally.

The relation of AMD with cataract was also examined for cataract subgroups based upon type of cataract. To be classified in either the "nuclear sclerosis" or "cortical cataract" subgroups, eyes were required to have cataracts identified in either the nuclear region or cortical region exclusively (as determined by a precoded variable in the data

TABLE 1—Frequencies of Persons with AMD and other Retinal Conditions by Lens Conditions, National Health and Nutrition Examination Survey, 1971–74

	AMD		NON-AMD			
			Other Retinal Abnormalities		No Retinal Abnormality	
	OD	OS	OD	OS	OD	OS
No Lens Opacifying Disease	57	61	676	670	1144	1125
Total Lens Opacifying Disease	103	95	614	638	493	498
Aphakics	6	6	21	25	25	26
Total Lens Opacities	97	89	593	613	468	472
Opacities without Visual Impairment	60	63	385	400	290	290
Total Cataracts	37	26	208	213	178	182
Nuclear	8	5	29	35	36	38
Cortical	5	3	58	53	55	54
Other	24	18	121	125	87	90
Total # Eyes	160	156	1290	1308	1637	1623

OD = oculus dexter (right eye)
OS = oculus sinister (left eye)

set). All other cataracts were placed in a category designated as "other" which included eyes with mixed nuclear and cortical cataracts.

Crude weighted and unweighted odds ratios were initially calculated for all age groups combined and then after stratification by 10-year age groups. The results for weighted and unweighted analyses were not appreciably different. Unweighted odds ratios based upon actual frequencies with confidence limits calculated using Woolf's method[16] are presented in the tables. Finally the relationship of AMD with lens opacities was examined using a logistic model controlling for age, sex, and a common suspected risk factor for both: systolic hypertension.[15] The model was run using both unweighted data and a supplementary logistic regression program in SAS designed to incorporate sampling weights and take the complex sample design into account.[17] Results were similar and presented for both weighted and unweighted data.

Results

The study population included 3,738 persons (6,174 eyes) (Table 1). The prevalence of both lens opacifying disease and AMD increased dramatically with age (Figure 1). Lens disease was generally bilateral. AMD was found in 185 persons (316 retinas). In 131 (70.8 per cent) subjects the disease was bilateral.

In the univariate analysis, persons with cataracts were substantially more likely to have AMD than were persons with no lens abnormalities (Table 2). Similarly, persons with lens opacities but no visual impairment had an odds ratio of 2.5 (95% CI = 1.8, 3.5) when compared with persons with no lens opacities. A person with at least one aphakic eye had an even higher odds ratio for AMD (OR = 4.6; 95% CI = 2.5, 8.6).

Estimates of odds ratios measuring the strength of association between AMD and three categories of lens abnormality (opacities without visual impairment, cataracts, aphakics) are presented for three age strata in Table 3. All estimates remain greater than 1.0, with lower bounds of the 95 per cent confidence limits excluding 1.0 in the highest age stratum (where most of the cases were found) for the opacity without visual impairment and aphakic groups. The data shown in Table 3 are for individuals, with a person defined as having the condition if it was observed in either eye. Similar results were observed when analyses were conducted for the left and right eyes separately. These estimates suggest that

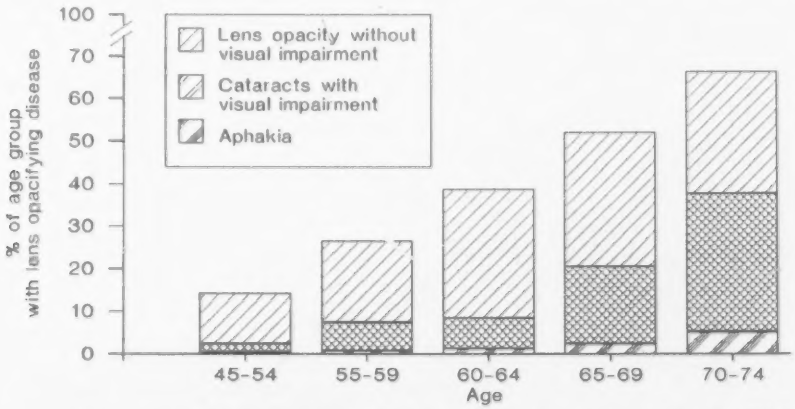

FIGURE 1a—Prevalence of Lens Opacifying Disease by Age Group in the Non-institutionalized Elderly Population of the Coterminus United States.
SOURCE: Data from the National Health and Nutrition Survey, 1971–74.

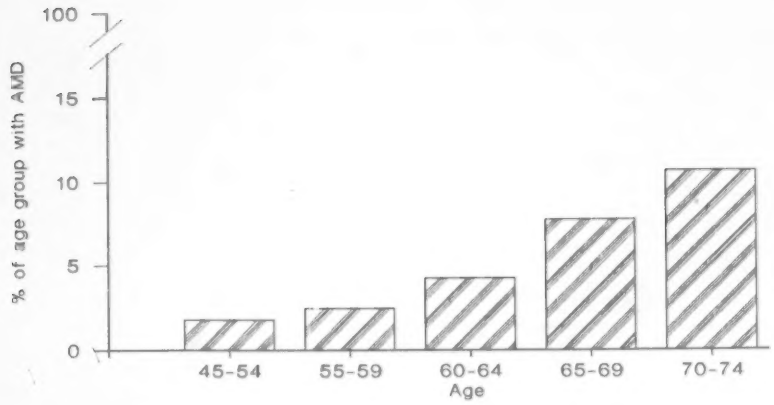

FIGURE 1b—Prevalence of Age-related Macular Degeneration by Age Group in the Non-institutionalized Elderly Population of the Coterminus United States.
SOURCE: Data from the National Health and Nutrition Survey, 1971–74.

TABLE 2—Crude Odds Ratios and 95% Confidence Intervals ( ) for Association of AMD and Various Lens Opacifying Conditions in the National Health and Nutrition Examination Survey, 1971–74

	OD	OS	Either eye
No Lens Opacity	1.0	1.0	1.0
Lens Opacity without Visual Impairment	2.8 (2.0, 4.1)	2.7 (1.9, 3.8)	2.5 (1.8, 3.5)
Cataract	3.1 (2.0, 4.7)	2.0 (1.3, 3.2)	2.4 (1.6, 3.7)
Nuclear Sclerosis	2.5 (1.6, 4.1)	2.4 (1.5, 3.8)	2.2 (1.4, 3.5)
Cortical Cataract	2.8 (1.8, 4.4)	2.3 (1.5, 3.6)	2.4 (1.6, 3.6)
Other Cataract	3.7 (2.3, 6.1)	2.5 (1.4, 4.2)	2.9 (1.9, 4.7)
Aphakia	4.2 (1.8, 9.5)	3.5 (1.4, 8.4)	4.6 (2.5, 8.6)

OD = Oculus dexter (right eye)
OS = Oculus sinister (left eye)

the odds for having AMD may be increased 50 percent among persons with cataracts, 80 percent among persons with non-visually impairing opacities, and 200 percent among aphakic persons.

Multiple logistic regression analyses were conducted to appraise the association between lens abnormalities and AMD while simultaneously taking into account other major predictors of macular degeneration. The effects of each of the three primary categories of lens disease were simultaneously tested in a model containing age, sex, and systolic blood pressure. As shown in Table 4, odds ratios determined by this approach were similar to those described above with age stratification. Again, the strongest effect was associated with aphakia which had a weighted odds ratio of 2.0, followed by non-visually impairing lens opacities with a weighted odds ratio of 1.8. These results suggest that the increment in odds ratio associated with aphakia (as an example) is similar to that associated with 10 years of additional age.

The number of persons with exclusively nuclear or

TABLE 3—Age-Stratified Odds of Association between AMD and Lens-Opacifying Disease, National Health and Nutrition Examination Survey, 1971–74

	Lens Opacity without Visual Impairment		
Age (years)	Cataract	(LO)	Aphakia
45–54	*	3.0 (0.9, 9.8)	*
55–64	1.5 (0.3, 7.1)	1.4 (0.5, 3.8)	*
65–74	1.5 (1.0, 2.4)	1.8 (1.2, 2.8)	3.2 (1.5, 6.5)
Mantel-Haenzel Age Adjusted Odds Ratio with 95% Confidence Intervals	1.5 (1.0, 2.3)	1.8 (1.3, 2.6)	3.3 (1.7, 6.5)

*insufficient data.
95% CI in parentheses.

TABLE 4—Weighted and Unweighted Multivariate Logistic Regression of AMD on Lens Opacifying Disease, Age, Sex, and Systolic Blood Pressure in NHANES I Elderly Population, 1971–74

	Weighted OR	Unweighted OR
Lens Opacifying Disease		
None	1.00	1.00
Opacity without Visual Impairment	1.80 (1.40, 2.30)	1.75 (1.22, 2.51)
Cataract	1.14 (0.84, 1.55)	1.29 (0.85, 1.95)
Aphakia	2.00 (1.44, 2.78)	2.46 (1.24, 4.88)
Age (per year)	1.07 (1.05, 1.10)	1.08 (1.05, 1.11)
Sex		
Male	1.00	1.00
Female	0.73 (0.54, 0.99)	0.76 (0.56, 1.03)
Systolic BP (per 10 mm Hg)	1.08 (1.04, 1.13)	1.06 (1.00, 1.12)

95% CI in parentheses.

cortical cataracts was small in this study, making detailed stratified analysis impossible. However, the crude odds ratios relating AMD to both cataract subtypes were strong and statistically significant. Neither the strength nor direction of the relationships differed, as has been suggested in other studies.[11]

Discussion

The analyses presented here suggest an association between two age-dependent eye abnormalities: age-related macular degeneration, and opacifying lens disease. The finding suggests that the pathogeneses of these two conditions may involve a common etiologic (exposure) factor, a common susceptibility element, and/or common disease-modulating risk factors. Since the pathogenesis of neither condition is well understood, we can only speculate about the nature of the common factor(s).

Development of lens and retinal diseases might reflect the independent or combined effects of such factors as: excessive light exposure, especially over a period of years; an inefficient constitutional anti-oxidant system such as the enzyme superoxide dismutase (SOD); the influence of deficient dietary antioxidants such as vitamins A, E, or C; and/or the damaging effects of oxidation-potentiating factors such as dietary psoralens or certain drugs.[18] These concepts suggest that some patients with concurrent lens and retinal disease might reasonably be viewed as suffering from accelerating aging, a formulation in line with prior speculations that cataracts may be a marker for advanced physiologic age.[19]

In the logistic regression, the presence of cataracts was weakly associated with AMD (OR = 1.1, p = 0.4), possibly reflecting the difficulty of visualizing maculas in eyes with dense cataracts, even with dilatation. It is appealing to theorize that the weak association between cataracts and AMD in this study may be due to the fact that cataracts retard the transmission of light to the retina, decreasing the extent of light damage. Unfortunately, limitations of the data regarding the severity of either AMD or cataracts preclude investigating this possibility more completely. An alternative explanation of the weak relation involves the possibility that different kinds of cataracts may have differing pathogeneses. For some types, no common factors may be shared with macular degeneration.

The very strong association of AMD with aphakia suggests that a sudden increase in light transmittance after cataract removal may reinitiate and dramatically accelerate the progression toward frank macular degeneration. This speculation provides additional reason to support the trend toward routine implantation of intraocular lens following cataract extraction, particularly with a lens capable of reducing the transmittance of ultraviolet or blue light. Alternatively, it has been suggested that cataract extraction can produce retinal changes resembling AMD.* It has also been suggested that exposure to bright surgical lights at the time of cataract extraction may be responsible for damage to the macula.[20] Serial examinations comparing patients with intraocular lens implants and those without lens implants would help resolve this question.

By defining all retinas without frank AMD as controls, there is a high probability of misclassifying retinas with subclinical AMD in the control group. The conservative nature of this approach enhances the believability of the association of AMD and lens opacifying disease which was found. When the analyses presented above were conducted using disease-free retinas as a control group, the relationships between AMD and all lens opacifying disease were stronger (higher estimates and tighter confidence limits).

The association between lens and macular disease independent of age is consistent with the theory of ultraviolet/near-ultraviolet light damage to the eyes. Until the actual risk factors for both conditions are clearly known, the most important implication of these findings is perhaps that surgical removal of cataracts, while improving vision in the short term, may actually place patients at greater risk of irreversible visual loss in the long run. With the increased life expectancy of the general population, this may become a serious problem, and warrants further investigation. The cheapest and most effective public health measure may be to encourage use of protective eyewear at early ages.

*Sperduto R: Personal communication, 1988.

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Five new pamphlets designed to help people caring for someone with Alzheimer's Disease (AD) have been developed by the Alzheimer's Association of Greater Washington. Alzheimer's families face unique challenges, since most care for their family member at home. The new pamphlets provide practical, up-to-date information to help families cope. Each pamphlet covers a specific Alzheimer's issue:

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- *Caregiving at Home* outlines the key practical issues involved in choosing to provide home care. It also provides many tips on safety, meals, personal care, leisure activities, and common behavioral problems.
- *Communicating with the Alzheimer Patient* examines how AD affects communication and explains many verbal and nonverbal techniques for helping the patient to communicate with and understand other people.
- *Financial Services You May Need* describes financial assistance and medical expense programs that can help a family meet the heavy cost of AD. It also tells what questions to ask, where and how to apply.
- *Directions in Alzheimer's Disease Research* discusses the latest findings about what causes AD, how it might be cured, what works to relieve symptoms, improving diagnostic methods and providing better care.

All the booklets indicate where to find more detailed information on specific issues. The set of five booklets is available from the Alzheimer's Disease Association of Greater Washington, 7970 Old Georgetown Road, Suite 4C, Bethesda, MD 20814. Include $2.00 for postage and handling.

Public Health Briefs

Self-Reported Diabetes in Mexican Americans: HHANES 1982–84

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Abstract: In the Hispanic Health and Nutrition Examination Survey (HHANES) of 3,928 Mexican Americans ages 20–74 years, the age-adjusted prevalence of self-reported diabetes was 6.8 percent among men and 7.6 percent among women. Comparable age-adjusted rates for the US population in a national survey were 2.9 percent in men and 3.8 percent in women. The prevalence of diabetes in Mexican Americans is greater in older age groups, was similar in men and women, and among women only was inversely associated with education. (*Am J Public Health* 1989; 79:770–772.)

Introduction

Community-based studies have shown that Mexican Americans in Texas have a three- to five-fold increase in the prevalence of diagnosed and undiagnosed non-insulin dependent diabetes mellitus when compared to non-Hispanic Whites.[1–4] The higher prevalence of obesity,[1–5] the more centralized fat distribution,[6] and genetic factors related to admixture with the indigenous peoples of the Americas among Mexican Americans may partly explain this increased prevalence.[3] Diabetes may not only be more prevalent but also more severe in its manifestations in this population.[7–9]

This report presents data on self-reported diabetes from the Hispanic Health and Nutrition Examination Survey (HHANES) and compares them to data from the second National Health and Nutrition Examination Survey (NHANES II).

Methods

HHANES was conducted in the southwestern United States (Arizona, California, Colorado, New Mexico, and Texas) between 1982–84. A stratified probability sample of the civilian, noninstitutionalized Mexican origin population was generated using a four-stage process.[10] The weighted estimates presented in this report are representative of Mexican Americans in those five states. Details of the HHANES sample design, operational plan, and quality control procedures have been published.[11]

The adult sample person questionnaire interviews were conducted by trained bilingual interviewers in the person's household in either English or Spanish. Mexican Americans ages 20–70 were asked: "Do you have diabetes or sugar diabetes?" Persons answering yes were then asked: "Did a doctor tell you that you have it?" Those who responded yes to both questions were defined as having self-reported diabetes.[12]

The acculturation scale developed for use in HHANES was based on eight of 20 items selected from an acculturation scale developed by Cuellar, *et al*:[13] an average score of eight items is based on language orientation (Spanish to English is scored from 1 to 5) and ethnic identification (Mexican to "American" scored from 1 to 5). For the purpose of this analysis, tertiles were used to categorize the acculturation scores into low, middle, and high groups.

Data on self-reported diabetes from the second National Health and Nutrition Examination Survey (NHANES II), a national probability sample of non-institutionalized civilians, are used for comparison.[14–16]

Statistical Analysis

Age-adjustment was carried out by means of direct standardization, using the US 1980 Census population as the standard.[17] Standard errors of age-adjusted rates were estimated using SESUDAAN.[18] For age-specific rates, the variances were calculated using the assumption of simple random sampling and then adjusted by the average design effect (1.0 in this case).[19,20] Association of self-reported diabetes with education, income, and acculturation were tested by multidimensional contingency table chi-square analysis, controlling for age and sex.

Results

A total of 4,735 Mexican Americans ages 20 through 74 were selected for the survey, with 3,928 (83 percent) completing the diabetes section of the household questionnaire. Respondents had a mean age of 36 years and were evenly divided between sexes; 64 percent were born in the United States. The comparison population from NHANES II consisted of 17,390 selected sample persons ages 20–74 years, of which 15,357 (88.3 percent) completed the diabetes questionnaire.

The age-adjusted prevalence rate for self-reported diabetes in Mexican Americans ages 20 through 74 years was 6.8 percent among men and 7.6 percent among women. Comparable rates for the US population measured in NHANES II were 2.9 percent for men and 3.8 percent for women.[17]

The prevalence of self-reported diabetes in Mexican Americans was greater in the older age groups of both sexes (Table

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1), as it was in the general US population.[17,18] There were no differences between rates in males and females within three of the four age groups of Mexican Americans, although prevalence in women 45 to 54 years of age was 1.8 times as high as in men. In comparison with NHANES II rates for each age group, self-reported diabetes was two to three times as prevalent for Mexican American men and women.

Increasing education among Mexican American women was associated with a significant decrease in age-adjusted rates of self-reported diabetes, but this difference was not noted among men (Table 1). The age-adjusted prevalence was higher in men with a middle tertile acculturation score (8.7 percent) than in those in the low (5.7 percent) or high (5.3 percent) tertiles. Family income was not associated with self-reported diabetes in either gender.

Self-reported treatment for diabetes was less common with insulin than with oral hypoglycemic agents (diabetes pills) in both Mexican Americans and the general population (Table 2). Adherence to special diets and to treatment with oral hypoglycemic agents was similar in the two populations. The proportion of Mexican American men who were aware of their diabetes, but not currently using any of these treatment alternatives, was considerably higher (40 percent) than reported by the NHANES II men with diabetes (27 percent).

Discussion

Because Mexican Americans are the fastest growing ethnic group in the southwestern states and diabetes is such an important contributor to morbidity and mortality in adults,[21] accurate information on prevalence is essential to public health authorities, clinicians, and community leaders.

The prevalence of self-reported diagnosed diabetes in adult Mexican Americans in HHANES is two to three times that estimated for White and Black adults in NHANES II and about twice the rate estimated for the general population from the 1982–84 National Health Interview Survey. Data from Bexar County, Texas, showed that compared with non-Hispanic

TABLE 1—Prevalence of Self-Reported Diabetes in Mexican Americans, Ages 20–74 Years, Mex.- an American Portion of HHANES, 1982

	Males (N = 1795)*			Females (N = 2133)*			Total (N = 3928)*		
	N	%	SE+	N	%	SE	N	%	SE
Age (years)									
20–44	1138	1.6	.38	1297	2.3	.41	2435	1.9	.28
45–54	323	6.1	1.33	421	11.0	1.53	744	8.7	1.03
55–64	231	17.7	2.51	273	17.3	2.29	504	17.5	1.59
65–74	103	24.2	4.22	142	21.4	3.44	245	22.7	2.68
Education (years)**									
0–6	571	7.8	.81	736	10.0	1.43	1307	9.0	1.08
7–11	459	5.9	1.09	589	6.1	.95	1048	6.0	.62
12+	724	7.8	1.73	765	3.9	.58	1489	6.0	.99
Income per Household ($$)**									
<$10,000	408	8.1	1.65	657	8.5	1.53	1065	8.3	1.00
10–19,999	604	6.1	.97	657	8.7	1.17	1261	7.4	.66
>20,000	621	8.1	2.81	611	6.2	1.64	1232	5.9	1.54
Acculturation**									
Low	550	5.7	.78	695	8.5	1.50	1245	7.3	1.07
Bicultural	675	8.7	.95	709	7.6	1.79	1384	8.1	1.02
High	530	5.3	2.03	672	7.8	1.76	1202	6.7	1.67

*Ns may not add up to totals because of missing data; data are weighted to the Mexican American population of the Southwest US.
**Age-adjusted to the 1980 US population
+SE = standard error.

TABLE 2—Treatment Regimens in Persons Ages 20–74 Years with Age-adjusted Self-reported Diabetes in HHANES and NHANES II

	HHANES (N = 257)*		NHANES II (N = 756)	
	Males (N = 107)	Females (N = 150)	Males (N = 318)	Females (N = 438)
Duration of diabetes (%)				
<5 yr	44	40	37	43
5–9 yr	22	22	25	22
10+ yr	34	38	37	35
Ever taken insulin (%)	32	42	35	34
Ever taken diabetes pills (%)	76	73	68	63
% Told to follow a special diabetes diet	85	92	73	76
% Currently following the diet	44	46	46	49
Current Therapy (%)				
Insulin	15	30	24	25
Insulin plus diabetes diet	10	16	17	16
Oral hypoglycemic agents	33	38	37	35
Oral agents plus diet	21	17	17	17
Specific diabetes diet alone	13	12	12	16
None of the above	40	19	27	25

*Numbers of sample persons are shown for information only; data are weighted to the Mexican American population of the Southwest US.

Whites, Mexican Americans have a higher mortality due to diabetes[22] and a greater prevalence of self-reported diabetes.[23] Rates for self-reported diabetes in Mexican Americans in Starr County, Texas,[23] and participating in the San Antonio Heart Study[2] (Unpublished data, Michael Stern) are similar for corresponding age groups to the HHANES rates for the entire southwest US Mexican American population.

The San Antonio Heart Study[2–4] has found that the age-adjusted diabetes rates for Mexican Americans are highest for those living in the Barrio (13.7 percent for men and 14.8 percent for women) and lowest among those living in the more affluent suburbs (6.1 percent for men and 3.7 percent for women).[4] Lower rates of diabetes among Mexican Americans with greater affluence and/or acculturation, however, were not consistently observed in HHANES. A significant inverse relationship between self-reported diabetes and education was found for women but not for men. Men with middle tertile acculturation scores reported a greater prevalence of diabetes, and no significant relationship was found for family income after adjusting for age. A possible explanation for the discrepancy in these two studies is that the San Antonio Heart Study results are based on total diabetes rates (diagnosed and undiagnosed) and the HHANES data in this report include self-reported diabetes only. When oral glucose tolerance tests have been used in population studies, the total prevalence of diabetes nearly doubles. Thus the socioeconomic gradient of less diabetes with increasing affluence may be present only if undiagnosed cases are included. Less affluent Mexican Americans may have a greater prevalence of undiagnosed diabetes because they have less access to health care.

The HHANES results in Mexican Americans confirm the presence of a two- to three-fold greater prevalence of self-reported diabetes, and health care providers need to give a high priority to diabetes as a major health issue in this group. Future research comparing Hispanic subgroups may help explain the relative importance of socioeconomic, cultural, and genetic factors of this observation.

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Black/White Differences in Non-treatment of Bladder Cancer Patients and Implications for Survival

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Abstract: Analysis of 20,764 White and 882 Black bladder cancer patients diagnosed during 1978–85 indicates that Black patients were more likely than White patients to go untreated following diagnosis after adjustment for age- and stage-at-diagnosis, sex, and tumor histology (OR = 1.80, 95% CI = 1.33, 2.43). Treatment status was found to be a significant predictor of five-year survival after adjustment (treated/untreated odds ratio = 3.16, 95% CI = 2.08, 4.79). Results suggest that differences in initial therapy may contribute to the survival differential between Black and White bladder cancer patients. (*Am J Public Health* 1989;79:772–774.)

Introduction

Decreased survival in Black patients relative to White patients with bladder cancer has been documented in several reports.[1–8] A number of factors have been found to contribute to this survival difference: age, diagnosis at more advanced stages, and more aggressive tumor histology. Within stage- and histology-specific groups, Blacks continued to have poorer five-year relative survival rates.[5,8] Findings by Axtell and Myers[7] suggest that different treatment patterns could also explain some of the racial differences in survival.

Questions remain regarding differences in treatment with respect to such determinants as age, stage at diagnosis, and histologic type, and their role in poorer survival among Black relative to White bladder cancer patients. We attempted to answer these questions through the study of a marker for treatment differences—non-treatment—and its association with survival in 20,764 White and 882 Black bladder cancer patients diagnosed between 1978 and 1985.

Methods

This study is based upon data collected by nine population-based tumor registries participating in the National

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Cancer Institute's SEER program (Atlanta, Connecticut, Detroit, Hawaii, Iowa, New Mexico, San Francisco-Oakland, Seattle-Puget Sound, Utah).[9] The variables in this study are determined in the SEER registries by review of hospital, clinic, private laboratory, private radiotherapy, private surgery, private practitioner (where the hospital record indicated therapy was to be received from the private practitioner), and nursing/convalescent home records. Case finding audits are conducted to ensure complete coverage of the SEER areas, and periodic training sessions and workshops are held to maintain quality control.

For this analysis, only patients with "Black" or "White" race noted in the medical record, and with newly diagnosed bladder cancer (ICD-0 188.0-188.9) as a single or first primary tumor were selected.[10] Excluded were 457 cases: patients diagnosed through death certificate or autopsy report alone (n = 49), patients with tumors not microscopically confirmed (n = 333), patients with unknown first course of therapy (n = 56), and patients with missing data for stage-at-diagnosis (n = 19). Using these criteria, 20,764 White and 882 Black bladder cancer patients were studied for racial differences in first course of therapy. Excluded cases had a somewhat higher proportion of patients who were older, Black, and female than did selected cases.* There were no other differences.

First course of therapy includes all cancer-directed therapy within four months of the initiation of therapy, which may occur at any time following diagnosis. A categorical, non-treated/treated variable was derived from the data. Black and White patients were then compared on this treatment status variable.

*Data available on request to author.

Survival time was defined as the period from the month of diagnosis as noted in the medical record to the earlier data of most current follow-up or December 1985, the most recent date for which virtually complete follow-up data were available.

Logistic regression analysis was used to explore the role of race in non-treatment of bladder cancer, adjusting for the potential confounding variables of age- and stage-at-diagnosis, sex, and tumor histology. This analysis was conducted for all study patients combined and separately by registry using the LOGIST procedure available through the SAS Institute.[12]

Because the proportional hazards assumption of a constant hazard ratio over time was not met by the data, logistic regression analysis was used to study the relation of treatment status with five-year survival with bladder cancer, simultaneously adjusting for race, age- and stage-at-diagnosis, sex, and tumor histology. This analysis was restricted to the 7,321 cases diagnosed between January 1978 and December 1980, and thus available for five years of follow-up. Of these patients, 4,254 were known to have survived for at least five years following diagnosis, and 3,067 were known to have died within five years of diagnosis. Cases diagnosed in this period alive at a most recent follow-up date less than five years after diagnosis were excluded from our analysis (n = 175). There was no difference between selected and excluded cases in the proportion of treated and untreated patients.

Results

Black and White, treated and untreated, bladder cancer patients are compared by age- and stage-at-diagnosis, sex, and tumor histology in Table 1. In the logistic regression (Table 2), Blacks were more likely than Whites to go untreated even with adjustment for age- and stage-at-diagnosis, sex, and tumor histology (O.R. = 1.80, 95% CI = 1.33, 2.43). A higher probability of non-treatment was also asso-

TABLE 1—Race and Treatment Status by Age, Stage,* and Histology

Patient Characteristics	Black Cases		White Cases		Untreated Cases		Treated Cases	
	Number	(%)	Number	(%)	Number	(%)	Number	(%)
Total	882	(100.0)	20,764	(100.0)	609	(100.0)	21,037	(100.0)
Age-at-diagnosis								
<45	51	(5.8)	982	(4.7)	24	(3.9)	1,009	(4.8)
45–54	104	(11.8)	1,936	(9.3)	28	(4.6)	2,012	(9.6)
55–64	228	(25.9)	4,927	(23.7)	112	(18.4)	5,043	(24.0)
65–74	286	(32.4)	6,466	(31.1)	151	(24.8)	6,601	(31.4)
75–84	166	(18.8)	4,719	(22.7)	178	(29.2)	4,707	(22.4)
85+	47	(5.3)	1,734	(8.4)	116	(19.1)	1,665	(7.9)
Sex								
Female	317	(35.9)	5,500	(26.5)	209	(34.3)	5,608	(26.7)
Male	565	(64.1)	15,264	(73.5)	400	(65.7)	15,429	(73.3)
Stage-at-diagnosis*								
0–I	492	(55.8)	15,153	(73.0)	261	(42.9)	15,384	(73.1)
II	31	(3.5)	545	(2.6)	1	(0.2)	575	(2.7)
III	137	(15.5)	2,361	(11.4)	35	(5.8)	2,463	(11.7)
IV	167	(18.9)	1,803	(8.7)	160	(26.3)	1,810	(8.6)
Unknown	55	(6.2)	902	(4.3)	152	(25.0)	805	(3.8)
Histology								
Papillary	15	(1.7)	645	(3.1)	15	(2.5)	645	(3.1)
Squamous	67	(7.6)	381	(1.8)	48	(7.9)	400	(1.9)
Transitional	356	(40.6)	6,910	(33.3)	272	(44.7)	6,996	(33.3)
Papillary-transitional	366	(41.5)	12,114	(58.3)	181	(29.7)	12,299	(58.5)
Other	76	(8.6)	714	(3.4)	93	(15.3)	697	(3.3)
Treatment status								
Treated	824	(93.4)	20,213	(97.4)	—		—	
Untreated	58	(6.6)	551	(2.7)	—		—	

*Data on the extent of disease at diagnosis were summarized to correspond with stages I-IV as described in the American Joint Committee's Manual for the Staging of Cancer.[11] These data include all information available within two months of diagnosis, excluding metastases known to have developed subsequent to diagnosis. *In situ* patients are included as stage I cases because their five-year survival rate is between the rates for stages I and II.

TABLE 2—Predictors of Non-treatment for Bladder Cancer: Analysis by Logistic Regression

Predictor Variable	Odds Ratio for Non-Treatment	95% CI
Race		
White	1.00	—
Black	1.80	1.33, 2.43
Sex		
Male	1.00	—
Female	1.18	0.99, 1.42
Stage		
0-I	1.00	—
II	0.08	0.01, 0.57
III	0.50	0.35, 0.73
IV	2.93	2.35, 3.66
Unknown	7.15	5.70, 8.97
Age-at-diagnosis*	1.03	1.02, 1.03
Histology		
Other	1.00	—
Papillary	0.33	0.19, 0.59
Squamous	0.97	0.66, 1.44
Transitional	0.46	0.35, 0.60
Papillary-transitional	0.22	0.16, 0.29

*per year of age.

TABLE 3—Predictors of Five-Year Survival with Bladder Cancer: Analysis by Logistic Regression

Predictor	Odds Ratio	95% CI
Untreated	1.00	—
Treated	3.16	2.08, 4.79
Race		
White	1.00	—
Black	0.54	0.40, 0.73
Sex		
Male	1.00	—
Female	1.33	1.17, 1.52
Stage		
0-I	1.00	—
II	0.42	0.30, 0.58
III	0.29	0.24, 0.34
IV	0.08	0.06, 0.11
Unknown	0.76	0.60, 0.96
Age-at-diagnosis*	0.92	0.91, 0.92
Histology		
Other	1.00	—
Papillary	3.03	1.98, 4.62
Squamous	0.70	0.42, 1.17
Transitional	1.71	1.26, 2.33
Papillary-transitional	3.07	2.26, 4.18

*per year of age.

ciated with advancing age, unknown stage and advanced stage at diagnosis. Patients with papillary, transitional cell, and papillary-transitional cell histology were less likely to be untreated as compared to patients with tumors histologically identified as squamous cell or "other" (including: "adenocarcinoma NOS," "malignant NOS" and "cancer NOS").

This logistic regression model was also used to analyze our data by registry. The non-treatment odds ratio for Blacks versus Whites was 1.74 (95% CI = 0.79, 3.82) in Atlanta, 3.10 (95% CI = 1.50, 6.41) in Connecticut, 1.38 (95% CI = 0.84, 2.26) in Metropolitan Detroit, 2.66 (95% CI = 0.29, 24.51) in New Mexico, and 1.60 (95% CI = 0.72, 3.56) in San Francisco-Oakland. The results were not interpretable for four other SEER registries due to small numbers and/or limited variable dispersion among the cases.

In a logistic regression analysis of survival, the odds ratio for five-year survival with bladder cancer for treated versus untreated patients was 3.16 (95% CI = 2.08, 4.79), with simultaneous adjustment for race, age- and stage-at-diagnosis, sex and tumor histology (Table 3). Analysis by registry yielded similar results in the eight registries for which there were sufficient data. Use of a proportional hazards model also yielded similar results.

The effect of treatment/non-treatment status on racial differences in survival was small (data not shown) as would be predicted given that non-treatment was selected as only a marker for treatment differences, with a relatively small proportion of patients in the non-treatment category.

Discussion

The results of this study indicate that there are differences in treatment status between Black and White bladder cancer patients, using as an indicator the proportion untreated. Black patients were more than twice as likely as White patients to go untreated. This difference was only partially explained by racial differences in age- and stage-at-diagnosis, sex, or tumor histology.

Non-treatment was also associated with poorer five-year survival among bladder cancer patients, even after adjustment for race, age- and stage-at-diagnosis, sex, and tumor histology. Race was also found to be associated with five-year survival with bladder cancer, even after adjustment for non-treatment. Given that treatment/non-treatment status was selected for use only as a marker for treatment differences, with a relatively small number of patients going untreated, we would not have predicted that adjustment for this variable would have a substantial impact on Black/White differences in survival.

These findings suggest that differences in treatment status contribute to the survival differential between Black and White bladder cancer patients.

Potential sources of bias in our findings include the possibility of race-selective underreporting of patient treatment. Because surgery and radiation therapy have been the mainstays of treatment for bladder cancer,[13] it is unlikely that initial therapy would occur outside settings where SEER reporting is virtually complete—hospitals, or private surgical or radiotherapy facilities. In addition, a conservative bias may have been introduced in the non-treatment odds ratio for women versus men by the exclusion of proportionately more women from the patient population selected for the study.

This study was limited by the type of data routinely collected by the tumor registries participating in the SEER program. SEER data on stage at diagnosis and tumor histology may not adequately adjust for these variables, particularly since we were unable to rule out bias in physician diagnostic practices as recorded in the medical record. SEER registries report data on populations in which rural Blacks are underrepresented. Thus, these results should be generalized to rural populations with caution. Treatment of cases was reported to SEER only in the broadest of categories (e.g., "surgery only"). As a result we were unable to assess the quality of care for patients who received some form of initial therapy. Individual patient characteristics which might be expected to influence treatment such as concomitant illness, education, income, and source of medical care could not be included in our analysis as they are not routinely reported to SEER.

A more complete assessment is needed of Black/White differences in patterns and quality of bladder cancer treat-

ment, their biologic, socioeconomic, and behavioral determinants, and their impact on survival. The National Cancer Institute is currently conducting a multicenter investigation of Black/White cancer survival differences that will further our understanding of these issues.

ACKNOWLEDGMENTS

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Factors Associated with Participation in a Community Senior Health Promotion Program: A Pilot Study

DAVID M. BUCHNER, MD, MPH, AND DAVID C. PEARSON, PHD

Abstract: Factors associated with participation in a community senior health promotion program were studied in 103 participants and a population-based control group of 531 non-participants. Compared to controls, participants had similar physical health status, but lower mental and social health status. Both men and women participants reported more depressive symptoms, lower positive affect, and lower social participation. Mental and social health may be important yet under-studied factors influencing participation in community health promotion programs. (*Am J Public Health* 1989; 79:775–777.)

Introduction

The results of large health promotion/disease prevention (HPDP) research projects[1–8] suggest HPDP programs attract relatively healthy persons in higher socioeconomic groups. But little is known about factors influencing participation in the community-based programs unaffiliated with a major research project—the setting in which the majority of health promotion/disease prevention programs presumably must occur. Also, few HPDP programs have studied recruitment of elderly subjects. For these reasons, we studied factors associated with participation in a community-based, senior health promotion program sponsored by a large health maintenance organization (HMO).

From the Department of Health Services and Center for Health Promotion in Older Adults, School of Public Health and Community Medicine, University of Washington, and the Health Services Research and Development Field Program, Seattle Veterans Hospital (Dr. Buchner) and the Center for Health Promotion, Group Health Cooperative of Puget Sound, Seattle (Dr. Pearson). Address reprint requests to David M. Buchner, MD, Department of Health Services SC-37, University of Washington, Seattle, WA 98195. This paper, submitted to the Journal March 7, 1988, was revised and accepted for publication August 15, 1988.

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Methods

The study was conducted at Group Health Cooperative of Puget Sound (GHC), a large, closed panel, not-for-profit HMO in western Washington State. The health promotion/disease prevention (HPDP) program (Growing Healthier) was intended for a broad target population of older adults and was advertised through the GHC magazine mailed to all enrollees, brochures distributed at GHC clinics, and presentations to consumer groups. The program was described as an opportunity to "enjoy life more" and "take greater control of your health and future." The curriculum consisted of a 10-week series of lectures, group discussions, and skills demonstrations led by trained instructors and senior volunteers. Specific topics covered included exercise, nutrition, stress management, social support, and self-responsibility/self-assertiveness.

Study participants were 103 (98 per cent) of the first 105 older adults (age 55+) to enroll in the Growing Healthier program given in the fall of 1984 at three of the 21 HMO clinics. Controls were 531 respondents (age 55+) to a survey of a stratified random sample of HMO enrollees (response rate = 90 per cent) and did not attend the program. For the analysis, control data were weighted to approximate a simple random sample.

The sources and/or definition of the independent variables used in this study are shown in Appendix I. Odds ratios assessed the association between program participation and subject characteristics. For consistency, variables with more than two levels were collapsed down to two categories. Adjustment for potential confounders was done using logistic regression.

Results

Table 1 describes the demographic characteristics of the study sample. Almost all study subjects were White. Compared to controls, participants were older, better educated, and reported higher incomes.

TABLE 1—Demographic Description of Study Participants and Control Group[a]

Variables	Men		Women	
	Participant N = 30	Control N = 235	Participant N = 73	Control N = 296
	%	%		
Age (years)				
55–64	23	52	34	49
65–74	63	35	55	34
75+	14	13	11	17
Race				
White	97	94	99	97
Other	3	6	1	3
Income				
<$15,000	23	41	32	54
$15000–25,000	15	32	32	26
>$25,000	62	27	36	20
Education				
<High School	13	32	8	30
High School	27	25	19	30
Some College	60	43	73	40
Marital Status				
Married	70	80	65	51
Other	30	20	35	49
Living Situation				
Lives Alone	21	16	30	39
with Spouse	76	79	68	52
Other	3	5	2	9

[a]Control data adjusted so that it reflects a simple, random sample of HMO enrollees over age 54 (actual sample was a random sample stratified by age).

Participants and controls differed on several study variables (Table 2). Use of seat belts and owning smoke alarms were associated with participation, but owning home fire extinguishers and refraining from drinking prior to driving were unrelated to participation. Lifestyle was associated with participation differently for men and women. Exercise and non-smoking were more strongly associated with participation in men. Obesity and abstinence from alcohol tended to be associated with participation for men, but with non-participation in women.

The health status of the participants and controls did not differ significantly as measured by self-perceived general health or by the Alameda Health Status scale (a measure of physical health). Varying the cut points on these scales in the statistical analysis did not change this result.

The mental and social health status of both men and women participants were lower than controls. Participation rates were substantially higher in persons with low positive affect, low emotional ties, high depressive symptoms, and low social participation. Low social support was not as strongly associated with participation. With the exception of emotional ties, findings were consistent between men and women.

Adjustment for education, income, marital status, and residential status had little effect on results, even though education in particular was associated with participation (age- and sex-adjusted odds ratio = 1.5, 95% confidence interval = 1.3, 1.8). This suggested the observed differences between participants and controls were not simply a reflection of the higher socioeconomic status of participants.

Discussion

Several findings of this study were unanticipated. First, associations of lifestyle factors with participation appeared to differ according to sex. Second, the general and physical

TABLE 2—Comparison of Participants in a Senior Health Promotion Program with Controls

	Men					Women					All[b]	
	Participant N = 30	Controls[a] N = 235	OR	Age-Adjusted OR	(95% CI)	Participant N = 73	Controls[a] N = 296	OR	Age-Adjusted OR	(95% CI)	Age Sex-Adjusted OR	(95% CI)
	%	%				%	%					
Lifestyle												
Non-smoker	97	79	7.5	6.3	(.95,48)	94	84	3.1	2.9	(.98,8.8)	3.6	(1.4,9.5)
Non-drinker	43	25	2.2	1.9	(.87,4.4)	25	36	0.6	.53	(.29,.98)		
Regular Exercise	50	18	4.5	4.4	(1.9,10)	31	23	1.5	1.5	(.76,2.8)	2.2	(1.3,3.6)
Overweight by >20%	26	17	1.7	1.8	(.70,4.7)	17	24	.65	.62	(.31,1.2)		
Home Car Safety												
Home Smoke Alarm	93	82	3.0	2.9	(.66,13)	95	83	3.5	3.2	(1.1,9.3)	3.1	(1.3,7.5)
Home Fire Extinguisher	69	67	1.1	1.2	(.51,2.7)	58	56	1.1	1.1	(.62,1.8)	1.1	(0.7,1.7)
Regular Seat Belt Use	73	46	3.2	3.0	(1.3,7.0)	66	49	2.0	1.9	(1.1,3.3)	2.2	(1.4,3.5)
Drinking Driver	23	27	.83	1.1	(.42,2.9)	5	7	.71	0.9	(.23,3.4)	.95	(.44,2.1)
Health Status–General												
Good/Excellent Health	83	78	1.4	1.5	(.56,4.3)	86	84	1.2	1.4	(.67,3.0)	1.4	(.80,2.6)
Health Status–Physical												
Low Alameda Score	83	72	2.0	1.7	(.63,4.8)	80	76	1.2	1.3	(.66,2.5)	1.4	(.81,2.4)
Heart Problems	13	14	.97	.88	(.28,2.7)	12	8	1.5	1.3	(.53,2.9)	1.1	(.55,2.1)
Diabetes	1	10	.14	.27	(.04,2.2)	7	11	.65	.58	(.21,1.5)	.40	(.16,1.04)
Emphysema	10	7	1.7	1.5	(.41,5.5)	6	6	1.1	1.1	(.33,3.2)	1.2	(.51,2.8)
Arthritis	43	30	1.8	1.8	(.83,4.0)	60	53	1.4	1.3	(.73,2.2)	1.4	(.91,2.3)
Health Status–Mental												
Low Positive Affect	44	26	2.3	2.5	(1.1,5.7)	41	27	1.9	1.9	(1.1,3.3)	2.1	(1.3,3.3)
High Depression Symptoms	43	21	2.7	2.6	(1.2,6.0)	45	27	2.2	2.1	(1.2,3.7)	2.3	(1.4,3.6)
Low Emotional Ties	52	50	1.1	1.2	(.57,2.7)	66	36	3.4	3.5	(2.0,6.2)		
Health Status–Social												
Low Social Support	36	28	1.4	1.5	(.64,3.5)	40	32	1.4	1.5	(.87,2.5)	1.5	(.93,2.3)
Low Social Participation	63	45	2.1	2.2	(1.02,4.9)	70	52	2.1	2.1	(1.1,3.7)	2.2	(1.3,3.4)

OR = Odds Ratio
95% CI = 95% Confidence Intervals
[a]Control data adjusted so that it reflects a simple, random sample of HMO enrollees over age 54 (actual sample was a random sample stratified by age).
[b]Summary odds ratio provided if strength of association similar for men and for women.

health of participants were no better than that of controls; although a lack of power or of sensitivity in the measurement scales could explain this result, in several measures of physical health the trends were for participants to be less healthy than controls. Third, participants had significantly lower mental and social health status, although not so low as to imply major social or psychological problems. As Growing Healthier targeted (in part) older adults wishing to improve their mental and social health, this finding is a notable exception to the usual situation of difficulty reaching a HPDP target population.

In contrast, participants in large HPDP research projects have often been physically healthier,[1,7] and the few available data suggest participants also tend to be socially and mentally healthier.[1,5,8] Perhaps factors which influence participation in large HPDP research projects differ from those in small community HPDP programs. Indeed, community programs lack several potential barriers to participation found in research studies (e.g., chance of randomization to a control group, burden of data collection, commitment to long-term follow-up).

But additional research is necessary to appraise the generalizability of these results. Factors influencing participation are to some extent program specific,[9] so some differences among programs are expected. As health status measures of this study were chosen because they were short and self-administered, research is needed using more comprehensive measures measuring a variety of aspects of mental and social health. Also, about three-fourths of participants received their medical care from the three clinics where the Growing Healthier program was given. Our data lacked clinic identifiers for participants, so we could not exclude the possibility that differences between participants and controls were due to differences between these three clinic sites and other HMO clinics. However, analysis of control data did not suggest these three clinics differed systematically from other HMO clinics on study variables.

In summary, mental and social health status may be important determinants of participation in community-based HPDP programs. Community HPDP programs should consider the potential importance of their mental and social health content. HPDP program evaluations should include social and mental health status measures. Of interest is whether program efficacy and effectiveness vary according to the mental and social health status of participants.

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APPENDIX I

Definitions of Independent Variables

Variables	Definition
Lifestyle	
Non-smoker	Not currently smoking cigarettes[10]
Non-drinker	No alcohol use in the past year[10]
Regular Exerciser	At least 20 minutes of exercise three times a week[10]
Overweight by >20%	Present weight greater than 120% of ideal body weight according to Metropolitan Life Tables
Home Car Safety	
Home Smoke Alarm	Smoke alarm(s) installed in the primary residence
Home Fire Extinguisher	Fire extinguisher present in the primary residence
Regular Seat Belt Use	Seat Belts usage reported as always or almost always (as opposed to never, rare, or sometimes)
Drinking Driver	Consumption of more than two drinks prior to driving an automobile in the past year[13]
Health Status–General	
Good/Excellent Health	Self-perceived health status was good or excellent (as opposed to fair or poor)
Health Status–Physical	
Low Alameda Score	Score of 1 to 4 on Alameda Health Status scale, indicating presence of at least one chronic medical condition (range of possible scores: 1 to 7)[12]
Heart Problems	Self-report of "heart trouble" in past year
Diabetes	Self-report of "diabetes" in past year
Emphysema	Self-report of "emphysema or chronic bronchitis" in past year
Arthritis	Self-report of "arthritis or rheumatism" in past year
Health Status–Mental	
Low Positive Affect	Score of 41 to 60 on Positive Affect Scale (ten questions, range of possible scores: 10 to 60)[11]
High Depression Symptoms	Score of 9 to 23 on Depression Scale (four questions, range of possible scores: 4 to 23)[11]
Low Emotional Ties	Score of 2 to 9 on Emotional Ties scale (two questions, range of possible scores: 2 to 12)[11]
Health Status–Social	
Low Social Support	Score of 4 to 7 on Social Support Scale (three questions, range of possible scores 1 to 7)[13]
Low Social Participation	Score of 2 to 3 on Social Participation Scale (two questions, range of possible scores: 0 to 3)[13]

New from nchs

Older People in the United States Who Receive Help with Basic Activities of Daily Living

MARY GRACE KOVAR, DRPH, GERRY HENDERSHOT, PHD, AND EVELYN MATHIS

There were about 28 million people in the United States who were age 65 or older in the mid-1980s. According to the US Bureau of the Census mid-level projection, there will be about twice as many, 59 million, by the year 2025.[1]

Two recent reports from the National Center for Health Statistics (NCHS), one based on nursing homes and their residents and the other on community-dwelling people, should help in planning for this increasingly large population of older persons. Together, these reports give a better profile of the population of elderly citizens who receive help with basic activities than has been available until now.

Data from the National Nursing Home Survey show that there were 1,318,000 individuals age 65 and older who were residents of nursing homes in 1985 (Table 1). Many of them were extremely old (45 percent were age 85 or older) because the rate of nursing home use is much higher among those who are age 85 and older than among the 65–74 or 75–84 age groups, as can be seen in Figure 1. Most of the nursing home residents required help with at least one basic activity and over half had difficulty with continence. Figure 2 compares the assistance required among the three age groups. The rates were consistently higher among those age 85 and older than among the younger residents.

The National Nursing Home Survey of nursing homes and their residents has been conducted three times: the first time was in 1973–74, the second in 1977, and most recently in 1985. The survey was based on a national (excluding Alaska and Hawaii) probability sample of nursing homes without regard to the level of care, whether they participated in Medicare or Medicaid, or whether they were licensed; board and care homes and residential care facilities were excluded.

Of 1,220 facilities selected for the sample, 57 were out of scope or out of business when the interviewer called; 84 did not participate. In the 1,079 participating facilities, five residents were selected from the list of residents in the facility on the night before the survey.[2] The member of the nursing staff most familiar with the patient's care answered the questions about the patient. Information was obtained for 97 percent of the 5,395 patients selected for this part of the survey.

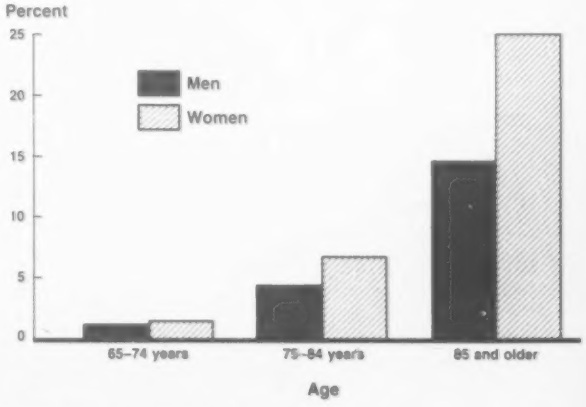
TABLE 1—Number of Persons Age 65 and Older in the United States

Age (Years)	In the Community 1984	In Nursing Homes 1985
	(In thousands)	
65 and older	26,433	1,318
65–74	16,288	212
75–84	8,249	509
85 and older	1,897	597

SOURCE: References 2 and 6.

However, the majority of persons who have passed their 65th birthday are not in nursing homes. In general, these individuals are younger than those in nursing homes; only 7 percent of the 26,433 people over age 65 living in the community in 1984 were age 85 or older. Data from the Supplement on Aging to the 1984 National Health Interview Survey show that about 2 million people, 7.6 percent of those age 65 and older who were living in communities, were receiving help with one or more of the specified activities shown in Figure 3.[3] The oldest group, age 85 years or older, who were living in the community were much more likely to be receiving help than community dwellers in their sixties or seventies and were more likely to be incontinent.

Descriptions of both the National Health Interview



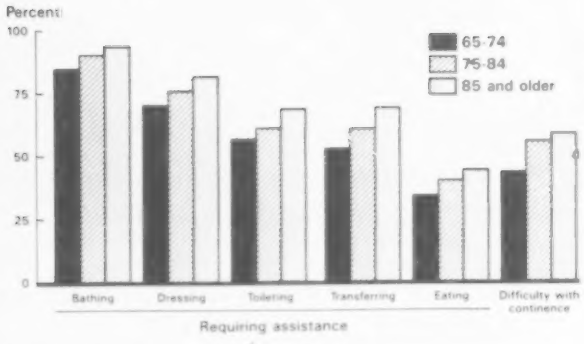
SOURCE: NCHS, National Nursing Home Survey, 1985.

FIGURE 1—Percent of US Population Age 65 and Older Who are Residents in Nursing Homes

Copies of "The National Nursing Home Survey: 1985 Summary for the United States" by E. Hing, E. Sekscenski, and G. Strahan *Vital and Health Statistics* Series 13, No. 97 are available from the US Government Printing Office, Washington, DC 20402. Order stock no. 017-022-01065-8.

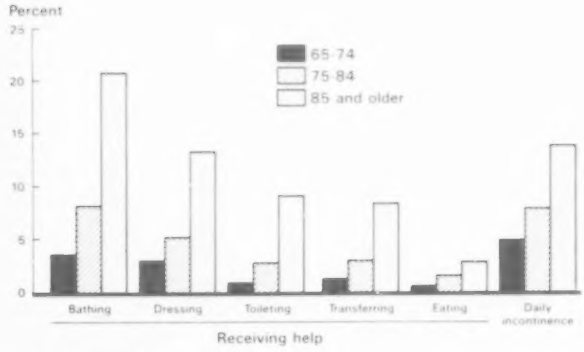
Copies of "Physical Functioning of the Aged: United States, 1984" by J. P. Fulton, S. S. Jack, S. Katz, and G. Hendershot, *Vital and Health Statistics* Series 10, No. 167 are available from the US Government Printing Office, Washington, DC 20402. Order stock no. 017-022-01069-1.

For more information about these and other reports and data from the National Center for Health Statistics, write to the Scientific and Technical Information Branch, NCHS, 3700 East-West Highway, Hyattsville, MD 20782, or call (301) 436-8500.



SOURCE: NCHS, National Nursing Home Survey, 1985.

FIGURE 2—Percent of Nursing Home Residents Who Require Assistance with Some Activities of Daily Living



SOURCE: NCHS, National Health Interview Survey Supplement on Aging, 1984.

FIGURE 3—Percent of Community Residents Age 65 and Older Who Receive Help with Some Activities of Daily Living

Survey (NHIS) and the Supplement on Aging (SOA) have been published by NCHS[4,5] and the NHIS has been described previously in this column. In brief, it is a continuing national survey of the health of people residing in households in the United States. In 1984, there were about 41,000 households in the sample; responses were obtained for 96.4 percent of them. The SOA was added that year to obtain information about older persons that had not been available previously from a national survey. All of the people age 65 and older and half of those ages 55–64 in responding households were included in the SOA sample; 96 percent participated in the SOA and 92 percent answered the questions for themselves.

The report from the National Nursing Home Survey[2] contains a great deal of additional information about the facilities, registered nurses, residents, and discharged patients. The estimated 19,100 nursing homes in the United States had 1,624,000 beds, and 794,000 full-time equivalent employees who provided direct health-related services to nursing home residents. There were 103,000 registered nurses, of whom 61,000 were full-time staff. The average per diem rate was $62 in Medicare certified facilities, $50 in Medicaid skilled, and $40 in Medicaid intermediate care facilities. For the 1,491,000 current nursing home residents, the median length of time they had been a resident was 614 days; for the 877,000 persons discharged alive, it was 70 days; for the 344,000 persons discharged dead it was 163 days.

Both residency in a nursing home and dependency in the community-living population are highly associated with age. The rate of nursing home residency was 18 times as high for people age 85 and over as for people ages 65–74. Comparable rates for receiving help with the specified activities among people living in the community were 5–6 times as high.

These high rates in very old aged are extremely important for planning because the age group defined as being age 85 and over is the age group whose size is increasing most rapidly. The number of people age 85 and over is projected to almost triple between 1985 and 2025—from 2,711,000 to 7,452,000 people.[1] Whereas they comprised 9.5 percent of the population age 65 and older in 1985, that figure is projected to be 12.7 percent in the year 2025.

The data from these reports, coupled with shortages of nurses, the relatively few children that people of the "baby-boom" generation have had, and the increasing involvement of women in the labor force, raise troubling questions about who will provide good care for an increasing number of dependent people if disability and dependency rates are not reduced through health promotion and disease prevention efforts and appropriate health care.

It is important to reduce these dependency rates because, if the current age-specific rates remain constant, the US could have 3.1 million persons residing in nursing homes in 2025, compared to 1.3 million in 1985; and we also could have 6.5 million people receiving help in one or more of the seven Activities of Daily Living for which information was collected on the Supplement on Aging compared to 2.5 million receiving such help in 1984.[6–8]

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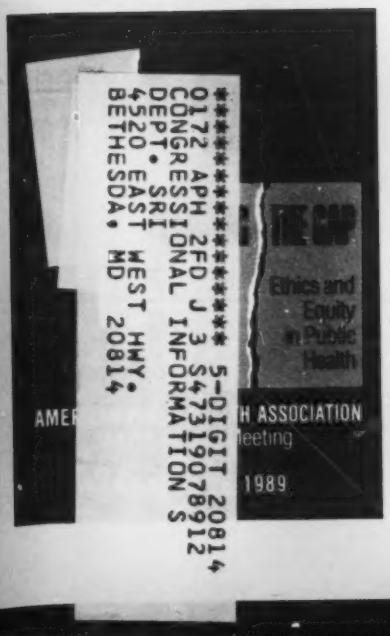

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Application of Clearance Concepts to the Assessment of Exposure to Lead in Drinking Water

FREDERIC Y. BOIS, PHARMD, THOMAS N. TOZER, PHD, LAUREN ZEISE, PHD, AND LESLIE Z. BENET, PHD

Abstract: This paper explores the application of clearance concepts to environmental toxicology. Lead, for which a clearance of about 0.5 ml/min is estimated from published data, is chosen as an example. An index for the contribution of drinking water to total exposure is developed using these concepts. For lead, this index is shown to increase with the concentration of the metal in water; it is higher for children than for adults. At the maximum contaminant level (MCL) of 10 μg/L proposed by the US Environmental Protection Agency (EPA), the average contribution from lead in drinking water is estimated to be 7 percent. The contribution in children is about twice as great. At and above the current MCL of 50 μg/L, drinking water becomes a major source of lead exposure. (Am J Public Health 1989, 79:827–831.)

Introduction

Regulatory standards, developed to protect the public's health, usually take neither variability in exposure nor sensitivity of subpopulations explicitly into account. Furthermore, the contribution from a particular contamination source (e.g., drinking water) is typically not evaluated with respect to its impact on the overall human exposure. A simple approach based on clearance concepts can provide an index of interindividual sensitivity, as well as an indication of the contribution to exposure by a particular source. Although clearance concepts are familiar to most clinicians, they are utilized to a much lesser extent for regulating environmental contaminants. The purpose of this paper is to explore this potential application.

The example for our purpose is lead. The common sources of lead are: drinking water, by corrosion of lead pipes and lead solders; air from the combustion of leaded gasoline; and soils, from the degradation of lead paints.[1–3] The toxicity of lead in humans has been extensively studied.[4–14] Even at low levels, cardiovascular changes in children's learning and behavior abnormalities are among several effects that are of concern. To limit human exposure, the US Environmental Protection Agency (EPA) has set a drinking water standard of 50 μg/L.[15] The same standard has been recommended by the World Health Organization.[16] This limit and the new value proposed by the EPA for lead in drinking water (10 μg/L at the tap, 5 μg/L entering the distribution system)[15] are subsequently evaluated with respect to the contribution of this source of lead to overall exposure using clearance concepts.

Clearance Concepts

The extent of exposure to pollutants can be assessed in terms of the blood concentration. There are many determinants of this concentration, including both absorption and elimination. Clearance is the most useful of the elimination parameters. Knowing the blood concentration under steady-state conditions, the daily intake can be estimated or, conversely, the blood concentration can be predicted when the rate of intake is known. The former use of clearance is emphasized in this paper.

Clearance, CL, is the proportionality constant between the rate at which the body eliminates an environmental toxicant, and its blood concentration, Cb. Under steady-state conditions the rate of elimination is equal to the rate of intake, Rt. Clearance is then a parameter (assuming first-order elimination) that relates blood concentration to the rate of intake; that is,

$$\text{Rate of intake} = Rt = CL \cdot Cb \qquad (1)$$

Consider a constant concentration, Cw, of lead in the water supply. The rate of intake of lead from drinking water, Rw, is the product of three terms: F, its availability or fraction absorbed into the general circulation, Qw, the water volume intake per day, and Cw. At steady-state, when water is the only source of exposure to the contaminant,

$$\text{Rate of intake} = Rw = F \cdot Qw \cdot Cw = CL \cdot Cb$$

or

$$Cb = (F \cdot Qw \cdot Cw)/CL \qquad (2)$$

Thus, under steady-state (or chronic) conditions the lead blood concentration is a function of water intake, lead concentration in the water supply, and lead blood clearance. Clearance is a property of both the administered compound and the individual receiving the compound.

Clearance concepts can be used to assess the contribution of a given source to the total exposure to a substance. Another use is to determine interindividual variability in both the exposure of handling of a toxicant. A third application is to aid in the determination of hazardous levels of toxicants. The first of these applications follows.

Given an estimated clearance value for a compound (CL_{ref}), measurements of blood concentrations in a local population, and drinking water concentrations to which the population is exposed, one can determine the fraction of total intake that can be attributed to drinking water. This fraction, Iw, is an index of contribution of drinking water to the input from all sources ($CL_{ref} \cdot Cb$):

$$Iw = Rw/Rt = (F \cdot Qw \cdot Cw) / (CL_{ref} \cdot Cb) \qquad (3)$$

This ratio ranges theoretically from 0 to 1.

Address reprint requests to Thomas N. Tozer, PhD, Professor of Pharmacy and Pharmaceutical Chemistry, University of California-San Francisco, School of Pharmacy, San Francisco, CA 94143-0446. Dr. Bois and Benet are also with the School of Pharmacy at UCSF; Dr. Zeise is with the California Department of Health Services/RCHAS, Annex 2, Berkeley, CA 94704. This paper, submitted to the Journal, February 27, 1988, was revised and accepted for publication October 20, 1988.

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Methods

Determination of Total Blood Lead Clearance

For analysis of isotopic tracer lead experiments, in which total input was controlled, equation 2 was used. When exposure to lead was not strictly controlled, the relationship between lead concentrations in blood and drinking water were determined by least-square regression, using a modification of equation 2. In this case, the rate of total lead intake Rt is comprised of two terms: a known exposure term Rw (through drinking water) and another term Rc accounting for concomitant (e.g., airborne, occupational, exposures.

According to equation 1:

$$\text{Rate of intake} = Rw + Rc = F \cdot Qw \cdot Cw + Rc = CL \cdot Cb$$

or

$$Cb = (F \cdot Qw/CL) \cdot Cw + Rc/CL \qquad (4)$$

Rc can be considered to be independent of Rw and the slope of the regression line of Cb on Cw gives CL/F.

Determination of the Index of Contamination

Equation 3 was used for all these determinations. The water and blood concentrations were obtained from the reports cited. Literature values[17] of water volume intake per day—2 L for adults (over the age of 15), 1.4 L for children, and 0.9 L for infants (under the age of 2)—were used. Values of 20 percent for adults and 40 percent for children and infants were assumed for the fraction absorbed F. These values are based on studies of the absorption of small amounts, less than 500 μg/day, of lead given orally in the diet. Absorption was on the order of 10 percent for normal, non-fasting adults[18] while, on an empty stomach, absorption can be 35 percent.[19] For children and infants, oral absorption has been observed to be higher, 40 to 50 percent.[2,20,21] These values can vary considerably from one individual to another.[20]

The reference clearance for children and infants has been obtained by scaling the value for a 70-kg adult using the empirical relationship:

$$CL_{ref}(\text{child}) = CL_{ref}(\text{adult}) \cdot [(\text{Body weight in kg})/70]^{0.7} \qquad (5)$$

This relationship corrects for differences in renal function under the assumption that lead clearance varies directly with renal function.[22]

The relationship between lead concentration in water and the index of contamination has been assessed by least-square regression. The difference between children and adults was tested by analysis of variance with a 5 percent significance level.

Results

Blood lead clearance and the contribution of drinking water to the total exposure of various populations, for which sufficient data are reported, are as follows.

Reference Value for Blood Clearance: Experimental Determination

Two sets of controlled experiments allow for a determination of reference values for blood lead clearance (see Table 1).

The first set of value is based on experiments previously performed.[23–29] Three male subjects were administered fixed amounts of lead in water drunk with each meal, for one to three years. Steady-state concentrations were reached. Data included the administered dose (300, 1000, or 2000 μg Pb/day), estimates of lead contained in the diet (derived from diet samples), as well as measurements of blood lead concentrations before and after exposure. The regression of blood lead concentration on daily intake gives a clearance estimate (inverse of the slope) of 0.7 ml/min (0.14 S.E.) assuming that, on average, 20 percent of the lead ingested is absorbed (i.e., F = 0.2). This calculation takes into account lead contributions from other sources, which were not controlled during the experiment itself.

TABLE 1—Blood Clearances for Lead Determined from Experimental Data of the Literature

Clearance (ml/min)	Remarks	Reference
0.7 ± 0.14	Subjects SW, MR, and EB	23–29
0.392	Subject — A	18,30
0.570	— B	—
0.324	— D	—

In the second set of experiments in human subjects,[18,30] ^{204}Pb nitrate was added in the diet. Experiments on three subjects lasted 104, 124, and 83 days, respectively. Although these exposures were of shorter duration than for the subjects of the first set of experiments, blood lead concentration-time profiles showed that steady-state conditions were nearly attained. Rather than supplementing the diet with lead, a fraction of the normal dietary intake of 300 μg/day was replaced with ^{204}Pb. Under these conditions, clearances were 0.392, 0.570, and 0.324 ml/min, respectively, for the three subjects with a mean value of 0.429 (coefficient of variation: 17 percent).

Lead Clearance Estimated from Epidemiological Surveys

Alternatively, lead clearance may be estimated from epidemiological studies in which individual lead concentrations in blood and drinking water are surveyed. Table 2 presents values obtained for various data sets using linear regression.

A regression performed on the first set of data[31] indicates a clearance of 0.44 ml/min (0.008 Standard Error) for males and 0.38 ml/min (0.008 SE) for females in a French population (correlation coefficient: 0.97). Exposures corresponded to lead concentrations in drinking water ranging from 0 to 400 μg/L, with a geometric mean of 160 μg/L.

Lacey, *et al*, analyzed the relationship between water lead concentration and blood lead in three-month-old infants.[32] The slope they determined corresponds to a clearance of 0.36 ml/min (0.07 S.E.); the correlation coefficient

TABLE 2—Blood Clearances for Lead Determined from Regressions of Published Epidemiological Data

Clearance (ml/min)	Standard Error (%)	N	Remarks	Reference
0.44	2	121	Males	31
0.38	2	132	Females	—
0.36	22	104	Infants	32
0.44	10	881	Males	33
3.8	NA	109	Females	34
1.8	NA	75	First draw water	35
1.2	NA	75	Running water	—

(NA: not available)

was not reported. Water lead concentration ranged from 0 to 600 μg/L. Total water intake was approximately 0.8 L/day, including water taken in infant formula. It is important here to correct for water intake and increased lead absorption in children (F = 0.4).

Blood and water lead levels in 24 British towns have been reported.[33] Mean water concentrations for the different towns ranged from 0 to 350 μg/L. In this concentration range the clearance is 0.44 ml/min (0.045 SE) and the correlation coefficient is 0.98.

A clearance of 3.8 ml/min is indicated in another source.[34] The standard deviation cannot be estimated from the data presented; it should be large, as indicated by a low correlation coefficient (0.36). It is also not certain how the points were weighted in the regression. Both these limitations apply to a further source of data[35] for which clearance values of 1.8 and 1.2 ml/min are estimated when first flush or running water samples, respectively, were used for assessing exposure. Correlation coefficients here were 0.76 and 0.72, respectively.

With the exception of the data from the earliest studies,[34,35] there is general agreement between the regression data and the values observed in controlled experiments. A blood clearance for lead of 0.5 ml/min for a 70-kg adult, with a range of 0.3 to 0.8 ml/min, therefore, seems reasonable. This value is used in the next section for determining the importance of lead in drinking water to an individual's overall exposure.

Contribution of Drinking Water to Total Exposure

The literature on lead measured in both an individual's blood and drinking water has been surveyed.[31,35–50] This survey is available from one of the authors (Tozer) upon request. These data allow estimation of *Iw* values.

The calculated index values range from 0.005 (0.5 percent) to 6.4 (640 percent) (values higher than 1 at high exposure may result from overestimation of the fraction absorbed or an underestimation of clearance). Figure 1 shows the frequency distribution of Iw; for 50 percent of the population, drinking water contributes at least 15 percent to the total exposure to lead, at least in our sample data. Figure 2 presents the relationship between Iw and the corresponding lead water concentration (in μg/L). Because of the apparent log-normal distribution of Iw and water lead concentration, the regression was performed using natural logarithms. When several measures of water lead concentration (e.g., first flush or running samples) were provided for the same population group, the median value was selected. The correlation coefficient for the regression was 0.90.

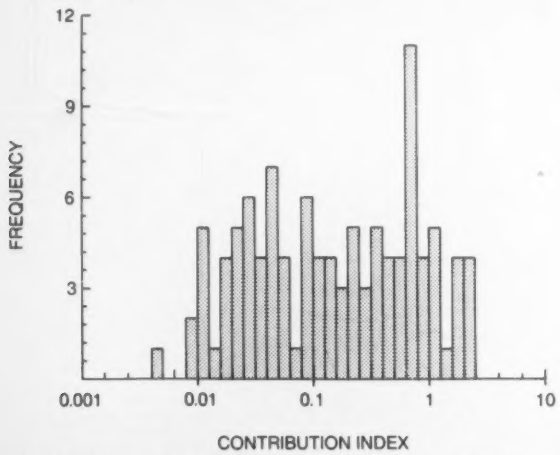

FIGURE 1—Distribution of the Index Iw of the Contribution of Lead in Drinking Water to Total Exposure, Determined from Epidemiological Data of the Literature (The Iw values are presented on a logarithmic scale).

Discussion

A clearance value obtained under steady-state conditions can be a useful indicator of the relative importance of chronic exposure to an environmental contaminant. Yet this concept has limitations and cannot be universally applied to all chemicals under all circumstances. As an example, consider compounds which are rapidly metabolized and/or excreted in the urine. A dynamic equilibrium is reached in which the blood concentration fluctuates dramatically over the day as a result of discontinuous intake (for drinking water) and high clearance. The proper assessment of a blood concentration would be difficult in this case. Lead, with its low clearance, does not fall into this category.

When a metabolite is primarily responsible for the toxicity, the use of the clearance of the parent compound as an indicator of individual sensitivity may be misleading. For example, a high clearance, for a given parent compound, could reflect a high rate of metabolite formation rather than a high renal excretion. This would imply an elevated target tissue dose, exactly contrary to what might be expected for the clearance value of the parent compound. Again, lead is not subject to this limitation.

Ideally the use of clearance concepts requires steady-state conditions and the presence of linear kinetics for distribution and disposition of the compound throughout the range of experimental and environmental exposures. For the majority of environmental contaminants, the assumption that levels of chronic exposure to a substance in air or water are

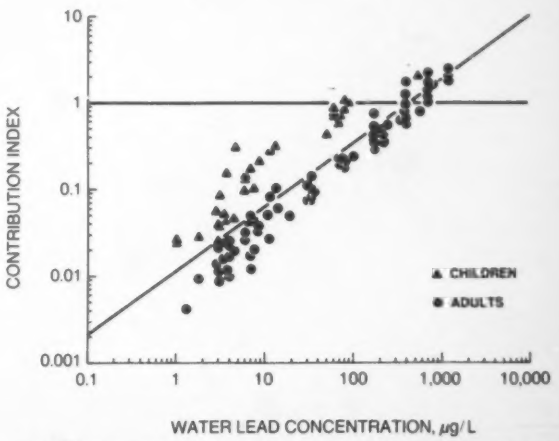

FIGURE 2—Regression of the logarithm of Iw on the logarithm of the lead concentration in water.
Data were obtained from published data in the literature. The regression equation is: ln (Iw) = −4.38+0.731·ln (lead concentration). The standard errors for slope and intercept are 0.0354 and 0.139, respectively. The correlation coefficient is 0.896.

below the values at which nonlinear behavior is observed is probably reasonable.

A nonlinear relationship seems to exist between blood lead and water lead concentrations at high exposures for humans.[33,36] Animal data indicate a possible saturation of blood lead levels.[52–54] Nevertheless, below approximately 500 μg/L the form of the relationship is nearly linear.[32] A clearance of 0.5 ml/min for adults is valid in this domain, and so are the results presented in Figures 1 and 2. Above 500 μg/L, Iw will be somewhat overestimated. A more mechanistic model[55–57] may more adequately describe the saturation of blood lead at high exposures. Two uptake mechanisms seem to be implied in intestinal lead absorption: a passive, non-saturable mechanism, and an active, saturable one.[5] A biphasic relationship between water lead and blood lead would be expected in this case. In this sense, the fact that the experimental clearance value determined in reference 23 (0.7 ml/min) is higher than that reported by Rabinowitz, *et al*,[18] could be explained by the much higher exposure in the former study.

Epidemiological studies may be used to estimate clearances through regression analysis. But high clearance values can be calculated if the intake of contaminated water is overestimated, for example by assuming a value of 2 L/day when the real value is much less. This situation may apply to the studies by Sherlock, *et al*.[49] A correct assessment of effective exposure is thus critical and a large population should be used.

Figure 2 shows that the contribution of lead in water to total lead exposure increases with its concentration at the tap. This relationship only occurs if the concentration of lead in drinking water correlates poorly with other sources of exposure and if drinking water is a major source of exposure. The lack of correlation of drinking water with other sources has also been reported.[50]

The spread about the regression line is quite large: the 95 percent confidence interval for estimates of Iw, at a given water concentration, span over a factor of 17. Therefore, for any particular water concentration the contribution of water to the total lead exposure cannot be calculated with accuracy. This variability may be explained by two factors. First, exposure is variable because water consumption is very different from user to user. We assumed standard values for different age groups.[17] Second, interindividual variability in lead absorption, distribution, and elimination clearly depends on individual factors.

It is not possible, in the absence of individual water consumption data, to separate these two sources of variation. To obtain a measure of interindividual variability, or identify different sensitive sub-groups, such information would be required. Nevertheless, in nearly all cases, the contribution of water to the overall exposure of lead is higher in children than in adults for the same water lead concentration. The fact that their intestinal absorption is higher by a factor of two contributes, together with their relatively higher water consumption, to this effect.

Note that for one-half of the population surveyed, waterborne lead represents less than 15 percent of the total exposure. The water for these populations contained only 30 μg/L or less of lead. When water lead concentration reaches 350 μg/L, on average 90 percent of total lead exposure arises from water. At the EPA MCL, 50 μg/L, about 20 percent of lead exposure in these populations is accounted for by lead in drinking water. However, at least 72 percent of lead exposure for 5 percent of the population might come from water contaminated with 50 μg/L of lead. This assumes a log-normal distribution (SE = 0.73) of *Iw* around the regression line.

We did not assume a dependence of lead clearance on urine flow. Under this hypothesis, the ratio Qw/CLref (in equation 3) would tend to have the same constant value for both adults and children. Such a flow dependence occurs when substances are extensively reabsorbed in the kidney tubules. Although lead seems to be partially reabsorbed,[58–61] renal clearance is not the only mechanism for lead elimination and the published data on renal lead handling in humans[62] suggest that lead clearance is not highly dependent on urine flow. To determine the impact of this flow hypothesis, we analyzed the effect of flow dependence. The difference between adults and children was reduced but still statistically significant.

ACKNOWLEDGMENTS

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ADDENDUM

Detailed lists of water lead concentrations and contribution index, Iw, of water to total exposure as determined from published epidemilogical data are available upon request to the authors.

Baby Doe Decision-Making in the 1990s: Conference Announced

A national conference entitled "Baby Doe Decision-Making in the 1990s" has been scheduled for September 14–15, 1989 at the Center for the Study of Bioethics, Medical College of Wisconsin, Milwaukee. Topics will include the effects of cost constraint measures, fetal therapy, genetic screening, and AIDS on Baby Doe decision-making.

The conference—co-sponsored by the Wisconsin Ethics Committee Network, the Medical Ethics Resource Network of Michigan, the Minnesota Network for Institutional Ethics Committees, and the Iowa Hospital Association—will be held at the Sheraton Mayfair Hotel in Milwaukee, WI.

For further information, contact: Anne Marie Talsky, Center for the Study of Bioethics, Medical College of Wisconsin, 8701 Watertown Plank Road, Milwaukee, WI 53226. Tel: 414/257-8498.

Survival Analysis of Hemophilia-Associated AIDS Cases in the US

JEANETTE K. STEHR-GREEN, MD, ROBERT C. HOLMAN, MS, AND MAUREEN A. MAHONEY, MT*

Abstract: Using national hemophilia-associated AIDS (acquired immunodeficiency syndrome) surveillance data and the life table method of survival analysis, the median length of survival of hemophilic patients in the United States after the diagnosis of AIDS was 11.7 months; the cumulative probability of survival at one year was 49.2 ± 2.0 percent; at two years, 28.9 ± 2.3 percent. Patients 13–29 years of age at the time of diagnosis had the longest survival and those 60 years and older had the shortest. Patients diagnosed since 1986 survived longer than those diagnosed before 1986. Length of survival did not differ significantly by race, coagulation disorder, AIDS manifestation at the time of diagnosis, or region of residence. Seven patients survived more than 36 months after AIDS was diagnosed. These patients were similar to those surviving for a shorter duration except that they were more likely to have met only the 1987 revision of the Centers for Disease Control AIDS surveillance case definition (as opposed to the 1985 case definition). Results of this study suggest that survival among hemophilic patients after the diagnosis of AIDS is similar to that reported by other investigators for other AIDS risk groups, excluding patients with Kaposi's sarcoma. (*Am J Public Health* 1989; 79:832–835.)

Introduction

The natural history of human immunodeficiency virus (HIV) infection in persons with hemophilia and other coagulation disorders is not well established and may differ from that in other acquired immunodeficiency syndrome (AIDS) risk groups.[1–3] Differences in the route and frequency of exposure to HIV, differences in the prevalence of potential cofactors for the development of AIDS (e.g., coexistent infections), and differences in routine medical care could lead to a variation in disease progression and ultimate outcome, such as survival. Using national surveillance data for hemophilia-associated AIDS, we examined the survival of hemophilic patients in the United States following the diagnosis of AIDS.

Methods

National surveillance of AIDS in persons with hemophilia and other coagulation disorders is maintained through the receipt of standardized AIDS case report forms submitted by individual state health departments to the Centers for Disease Control (CDC).[4] CDC personnel obtain supplemental information, including update of living status, through routine telephone interviews with the treating physician, nurse, or public health worker.

Only patients with illnesses that meet the CDC surveillance case definitions for AIDS, as revised in 1985 and 1987,[5,6] are included in the analyses. To be classified as having hemophilia-associated AIDS, the patient must have received blood or blood components since 1978 as treatment for a dysfunction in the normal coagulation system. Patients with multiple risk factors for HIV infection (i.e., risk factors in addition to a coagulation disorder, such as a history of male to male sexual contact or intravenous drug use) were excluded from the analyses (n = 24). Only patients reported before August 31, 1988 are included in the analyses.

The date of AIDS diagnosis was defined as the date the first AIDS-indicator disease (i.e., opportunistic infection or cancer meeting the CDC AIDS case definition)[5,6] was diagnosed. Survival time in months for each individual was calculated from the month of AIDS diagnosis to the month of death, if the patient had died, or from the month of AIDS diagnosis to August 1988, if the patient was still alive. Because length of survival was expressed in month intervals, the life table method of survival analysis was used to estimate the survivorship function from censored data.[7] The version of the Kruskal Wallis or generalized Wilcoxon test proposed by Breslow was used to compare the distribution of survival among patient subgroups.[8] CDC personnel obtained additional information on patients surviving longer than 36 months after the diagnosis of AIDS, including development of subsequent disease, other reasons for immunosuppression, and use of anti-HIV therapies, through telephone interview of the patient's physician, nurse, or associated public health worker in June 1988. Median ages at diagnosis of these patients and those surviving shorter periods were compared using Wilcoxon rank-sum test.[9] The method of Thomas and Gart was used to calculate odds ratios and 95% confidence intervals for dichotomous variables.[10]

Results

Between January 1, 1981 and August 31, 1988, 716 cases of hemophilia-associated AIDS (with no other risk factors for HIV infection) were reported to CDC. Six hundred and twenty-eight (88 percent) patients had hemophilia A (factor VIII deficiency); 61 (9 percent) had hemophilia B (factor IX deficiency); 16 (2 percent), von Willebrand's disease; five (<1 percent), acquired inhibitor to factor VIII; two (<1 percent), factor VII deficiency; two (<1 percent), factor X deficiency; and one each factor V and factor XI deficiency. All except 18 (3 percent) patients were male. The median age at the time of AIDS diagnosis was 31 years; however, patients diagnosed before 1986 were significantly older than those diagnosed in 1986 or later (median age of 34 years vs 30 years, $p < 0.001$). Of the 716 patients, 611 (86 percent) were White; 46 (6 percent), Black; 47 (7 percent), Hispanic; nine (1 percent), Asian/Pacific Islander; and three (<1 percent), American Indian/Alaskan Native.

Living status and date of death were known for 714 patients. As determined by the life table method, median survival from the time of AIDS diagnosis was 11.7 months. The cumulative probability of survival at one year was 49.2 ± 2.0 percent and at two years was 28.9 ± 2.3 percent.

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Survival did not differ by race, coagulation disorder, or region of residence (Table 1). Patients 13–29 years of age at the time of AIDS diagnosis had a significantly longer survival than the older patients 30–59 years of age and ≥60 years of age (Table 1 and Figure 1). Patients <13 years of age and 30–59 years of age also had a significantly longer survival than patients 60 years and older.

Patients diagnosed with AIDS before 1986 had a shorter survival than those diagnosed in 1986 or later (Table 2). Because the number of patients in our study diagnosed with only one AIDS-indicator disease other than *Pneumocystis carinii* pneumonia or esophageal candidiasis was small, survival by AIDS-indicator disease was limited to these two diagnoses. No difference in survival was observed (Table 2). Survival did not differ by case definition used to diagnose AIDS (1985 vs 1987 revision of the CDC AIDS case definition) (Table 2). Because the median age at diagnosis was younger among patients diagnosed in 1986 or later, analysis of survival time by year of diagnosis was stratified by age group (Table 3). In each age group, patients diagnosed since 1986 had a longer survival.

Seven patients survived longer than 36 months after the diagnosis of AIDS (Table 4). All were male; six had hemophilia A. These patients were similar to the patients surviving for shorter periods with respect to age at AIDS diagnosis, sex, race, and coagulation disorder. They were more likely to have met only the 1987 case definition (diagnosed retrospectively) as opposed to the 1985 case definition (4/7 [57 percent] vs 122/709 [17 percent], odds ratio = 6.4, 95% CI = 1.1, 44.2). All seven had antibody to HIV. Only one had been cultured for HIV and was found to be positive. No strain typing had been attempted on any of the patients. Five of the seven had been treated with some form of anti-HIV therapy; unfortunately, similar data were not available for patients living for a shorter duration.

Discussion

Our study, including patients with hemophilia and other coagulation disorders exclusively, expands upon studies of the natural history of HIV infection among different groups

TABLE 1—Length of Survival after Diagnosis of AIDS among Hemophilic Patients by Patient Characteristics, United States, January 1981–August 1988

Characteristics	Number	Median Survival Estimate (in months)	Cumulative Probability of Survival at: One Year	Two Years
Total	714	11.7	49.2 ± 2.0%	28.9 ± 2.3%
Age Group* (years)				
<13	60	12.0	50.1 ± 6.6%	34.8 ± 7.0%
13–29	255	16.4	60.2 ± 3.3%	40.3 ± 4.1%
30–59	333	9.8	44.5 ± 2.9%	23.3 ± 3.2%
≥60	66	3.0	28.7 ± 6.3%	5.4 ± 4.9%
Race				
White	611	11.9	49.5 ± 2.1%	28.7 ± 2.5%
Black or Hispanic	92	10.5	45.3 ± 5.9%	26.2 ± 6.4%
Coagulation Disorder				
Hemophilia A	626	11.4	48.1 ± 2.1%	27.5 ± 2.5%
Hemophilia B	61	14.9	60.7 ± 6.5%	42.7 ± 7.8%
Region of Residence				
Northeast	174	13.8	53.0 ± 4.0%	31.1 ± 4.7%
Northcentral	156	9.7	45.2 ± 4.1%	30.8 ± 4.3%
South	188	11.4	47.6 ± 3.9%	25.5 ± 4.8%
West	189	12.1	50.3 ± 4.0%	24.7 ± 5.3%

*$p < 0.005$ using the Breslow test.[8]

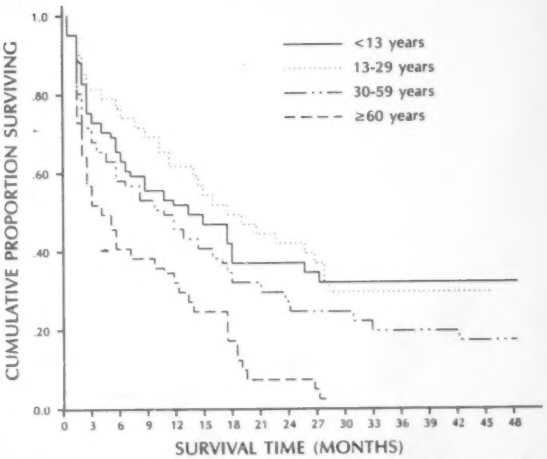

FIGURE 1—Survival after the Diagnosis of AIDS among Hemophilic Patients by Age at Diagnosis, United States, January 1981–August 1988.

TABLE 2—Length of Survival after Diagnosis of AIDS among Hemophilic Patients by Patient Characteristic, United States, January 1981–August 1988

Characteristics	Number	Median Survival Estimate (in months)	Cumulative Probability of Survival at: One Year	Two Years
Year of Diagnosis*				
Before 1986	185	6.4	36.2 ± 3.5%	14.6 ± 2.6%
1986 and after	529	15.0	54.6 ± 2.3%	39.4 ± 3.2%
AIDS Diagnosis				
P. carinii pneumonia only	302	14.2	54.2 ± 3.0%	29.9 ± 4.0%
Esophageal candidiasis only	57	16.0	55.7 ± 7.2%	37.6 ± 9.0%
Case Definition				
1985/1987	593	11.5	48.4 ± 2.2%	27.3 ± 2.4%
1987 only	121	13.7	54.8 ± 4.9%	44.9 ± 6.2%

*$p < 0.005$ using the Breslow test.[8]

TABLE 3—Median Length of Survival by Year of Diagnosis and Age Group

Age Group (years)	Year of Diagnosis: Before 1986 N	Before 1986 Median Survival Time (in months)	1986 and after N	1986 and after Median Survival Time (in months)
<13	14	—	46	10.9
13–29*	50	11.0	205	>29.0
30–59*	103	6.3	230	11.9
60+	18	1.7	48	4.9

*$p < 0.005$ using the Breslow test.[8]

at risk for AIDS. Rothenberg, *et al*, analyzed data from 5,833 AIDS patients in New York City.[11] Median length of survival ranged from 9.4 months in patients with a history of intravenous drug use to 13.1 months in patients with a history of homosexual activity. Patients with both risk factors had a median survival of 10.1 months and those with other risk

TABLE 4—Seven Patients with Hemophilia-Associated AIDS Surviving Longer than 36 Months after AIDS Diagnosis

Patient	Age at Diagnosis	Date of AIDS Diagnosis	AIDS Indicator Disease	Comment
A	35	1/84	cryptococcal meningitis*	on prednisone for thrombocytopenia at time of meningitis; treated with AZT and acyclovir; still living**
B	48	1/85	presumptive esophageal* candidiasis	treated with AZT; subsequently developed HIV encephalopathy; still living
C	10	9/83	recurrent bacterial infections*	treated with AZT; subsequently developed a lymphoma; still living
D	36	11/84	presumptive esophageal candidiasis*	treated with AZT; eventually died with severe liver disease
E	31	3/84	Kaposi's sarcoma	treated with radiation therapy; still living
F	37	3/84	*P. carinii* pneumonia	treated with gamma interferon and ribavirin; still living
G	13	12/84	*P. carinii* pneumonia	subsequently developed esophageal candidiasis, cryptosporidiosis, and *Mycobacterium avium-intracellulare;* treated with AZT; still living

*met 1987 CDC AIDS surveillance case definition only.
**as of June 1988

factors had a median survival of 8.5 months. In an analysis of 663 AIDS cases from the United Kingdom,[12] Reeves and Overton reported a median survival of 12.6 months in AIDS patients with a history of homosexual activity and 10.5 months in other transmission categories combined. Both studies found the most favorable survival among patients with Kaposi's sarcoma alone (25.0 and 15.0 months, respectively). Neither study was able to examine hemophilic patients separately. Our results suggest that survival among hemophilic patients with AIDS is comparable to that observed in other patient groups, excluding those with Kaposi's sarcoma. However, data from studies by Rothenberg and Reeves were collected in 1985 and 1987, respectively. Survival may have improved in these groups. In our study, survival improved over time. This improved survival could reflect changes in health care for AIDS patients including earlier diagnosis and institution of treatment of AIDS-indicator diseases, prophylactic treatment for some opportunistic infections (e.g., *Pneumocystis carinii* pneumonia),[13] and development of anti-HIV therapies such as azidothymidine.[14,15] To facilitate comparisons between patient groups, survival analyses should be repeated in other groups to determine the impact of more recent changes in health care.

Similar to studies by Rothenberg and Reeves, we found a decreasing length of survival after the diagnosis of AIDS with increasing patient age. It has been shown that antigenic stimulation causes increased replication of HIV which may lead to a more rapid development and progression of HIV-related disease.[16] Repeated exposure to factor concentrate (also a powerful antigenic stimulus),[17] exposure to environmental contaminants, and a higher prevalence of hepatitis B and sexually transmitted diseases (Stehr-Green JK: unpublished data) could lead to a more rapid progression of HIV infection among older hemophilic patients and a less favorable survival. A more rapid progression of infection could also explain the decrease in median age of hemophilia-associated AIDS cases over time, consistent with findings of Eyster, *et al,* which showed age to be a cofactor for AIDS independent of date of infection.[18]

The 1987 revision of the CDC AIDS surveillance case definition includes a broader range of AIDS-indicator diseases (including HIV dementia and HIV wasting syndrome), indicator diseases that are presumptively diagnosed, and elimination of exclusions due to other causes of immunodeficiency. The large proportion of long survivors with illnesses meeting only the 1987 revision seems to suggest that these illnesses may be associated with a more favorable survival than those included in the 1985 case definition. However, the illnesses included in the revision may allow the patient to be diagnosed with AIDS and reported to CDC at an earlier point in the progression of HIV-related disease. That is, the interval between infection and development of clinical manifestations of AIDS may be shortened, resulting in an apparent increase in the length of survival after the diagnosis of AIDS. Analyses of survival from date of infection to death could eliminate this potential confounder. Unfortunately, the date of seroconversion is not available for most hemophilic patients due to multiple exposures to potentially contaminated blood products.

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$5 Million Awarded to States to Help Low-Income AIDS Patients Purchase Medicines

An additional $5,075,490 in federal funds from the Health Resources and Services Administration (HRSA) of the US Public Health Service was recently awarded to 26 states to help low-income AIDS patients obtain drugs that would prolong their lives. The Food and Drug Administration (FDA) has determined that three drugs prolong the lives of AIDS patients—azidothymidine (AZT), alpha interferon, and aerosolized pentamidine.

The funds were distributed on March 31 just before the authorizing legislation, Public Law 100-471, expired. Those eligible for assistance are low-income individuals not covered by Medicaid or private insurance. Each state determines eligibility. The 26 states receiving the funds are: Alabama, Alaska, Arizona, California, Colorado, Florida, Georgia, Idaho, Illinois, Indiana, Kansas, Kentucky, Maine, Massachusetts, Mississippi, North Dakota, Ohio, Oklahoma, Oregon, Rhode Island, South Carolina, Tennessee, Texas, Utah, Vermont, and Wisconsin.

Of the total distributed, $5 million came from the transfer of fiscal 1989 funds appropriated for AIDS programs administered by the Office of the Assistant Secretary for Health and four Public Health Service agencies. The remaining $75,490 derived from jurisdictions that reported they would not be able to use by September 30, 1989 funds distributed last November under PL 100-471. At that time, HRSA distributed to the 50 states and other jurisdictions some $15 million including $5 million from a private pharmaceutical company. The drug reimbursement program began in fiscal 1987 when HRSA awarded $30 million to states. The latest distribution brings the total federal funds awarded under the AIDS Drug Program to $45 million.

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Birthplace and the Risk of AIDS among Hispanics in the United States

Richard M. Selik, MD, Kenneth G. Castro, MD, Marguerite Pappaioanou, DVM, PhD, and James W. Buehler, MD

Abstract: To extend previous work showing that the risk of AIDS (acquired immunodeficiency syndrome) is higher in US Hispanics than in Whites who are not Hispanic, we compared US residents born in different Latin American countries. We computed the cumulative incidence (CI) of AIDS and the distribution of cases by mode of exposure. Cases were those reported to the Centers for Disease Control between June 1, 1981 and December 12, 1988, and populations specific for birthplace were from the 1980 census. The reference group was the White population that was not Hispanic, CI 25.7/100,000. We estimated a similar rate in Mexican-born persons (25.3/100,000). In the South and West, the rate in Mexican-born Hispanics was half the reference rate. In each US region, the CI of AIDS in heterosexual intravenous drug abusers (IVDAs) in Puerto Rican-born persons was several times greater than that in other Latin American-born persons. Puerto Rican-born persons were the only Latin American-born persons in whom most cases were in heterosexual IVDAs. The data suggest that resources for preventing AIDS in Hispanics are needed most in those of Puerto Rican ethnicity for AIDS related to intravenous drug abuse. (*Am J Public Health* 1989; 79:836–839.)

Introduction

The risk of acquired immunodeficiency syndrome (AIDS) is higher in US Hispanics than in Whites who are not Hispanic.[1,2] Since Hispanics are a diverse ethnic group, targeting public health resources for preventing AIDS in Hispanics would be improved by information on AIDS risks in different Hispanic ethnic subgroups. Because data on AIDS cases reported to the Centers for Disease Control (CDC) include country of birth but not country of ancestry, we investigated the variation in the risk of AIDS by country of birth.

Methods

We analyzed the 78,751 AIDS cases meeting the case definition[3] reported to CDC between June 1, 1981 and December 12, 1988 in residents of the United States (the 50 states and District of Columbia).* We classified persons with AIDS according to their race/ethnicity and country or territory of birth. We computed the cumulative incidence, the number of cases per 100,000 US population of the respective group (1980 US census),[4] and the relative risk (RR), the ratio of the cumulative incidence in a group to the cumulative incidence in Whites who were not Hispanic. The 99% confidence limits around RRs were approximated by the first-order Taylor series method.[5] Since census data were unavailable for birthplace by race/ethnicity, we estimated the population of US-born Hispanics as the difference between resident Hispanics and residents born in Latin America.

Analyses were stratified by geographic region (Northeast, Midwest, South, and West)[4] and by presumed mode of acquiring human immunodeficiency virus (HIV) infection.[6] To detect the effect of differences in the prevalence of the various means of acquiring HIV in different populations, we stratified only the AIDS patients by exposure category, not the populations.

Results

Of the 10,336 Hispanic persons with AIDS in the United States, 49 percent were reported to have been born in the United States, 37 percent in a specified Latin American country or territory, 1 percent in another specified country, and 4 percent in an unspecified foreign country; the birthplace of the remaining 10 percent was unreported. Of the 4,045 Latin American-born persons with AIDS in the United States, 94 percent were Hispanic.

The cumulative incidence of AIDS in all Hispanics was elevated (70.8/100,000), but was not elevated in persons born in Mexico (25.3/100,000) (Table 1). In the South and West, the cumulative incidence in Mexican-born persons and in US-born Hispanics (≥75 percent of whom were of Mexican ethnicity in those regions) were both lower than the reference rate, but in the Northeast and Midwest they were higher than the reference rate.

The pattern of risk by ethnicity was similar for AIDS in homosexual/bisexual men without a history of intravenous drug abuse, representing 61 percent of all cases. In addition, the cumulative incidence in US-born Hispanics was about the same as the reference rate in the nation as a whole and in the Midwest (where 70 percent of US-born Hispanics were of Mexican ethnicity). A similar pattern was seen for AIDS in homosexual/bisexual male intravenous drug abusers (IVDAs), representing 7 percent of all cases; the cumulative incidence in Mexican-born persons was less than the referent in the West and the nation as a whole, and similar to the referent in the other three regions.

For AIDS in heterosexual IVDAs, representing 19 percent of all cases, the cumulative incidence in Mexican-born persons was above the referent in the Northeast and similar to the referent elsewhere (Table 1). The cumulative incidence in US-born Hispanics and persons born in Puerto Rico and Cuba exceeded the referent in each region, but among those born in other Latin American countries the cumulative incidence was greater in the Northeast region only. In each region, the cumulative incidence was markedly greater in Puerto Rican-born persons than in persons born in Cuba, Mexico, or other Latin American countries.

In the nation as a whole, Puerto Rican-born persons were the only Latin American-born persons in whom most AIDS cases were in heterosexual IVDAs (Table 2). Among persons born in Mexico, Cuba, and other Latin American countries, the proportion of cases in heterosexual IVDAs was 10 percent or less. Among the Puerto Rican-born, the proportion of cases in heterosexual IVDAs ranged from 32 percent in the South and West to 52 percent and 61 percent in the Midwest and Northeast, respectively—proportions

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*NOTE: Residents of Puerto Rico or other US territories were excluded.

TABLE 1—Cumulative Incidence (CI)* and Relative Risk (RR)† of AIDS in Hispanics and/or Latin American-Born Persons, by Ethnic Group§ and US Region, for Different Exposure Categories

Region	Ethnic Group§	Population (1980 Census)	Exposure Category of AIDS Cases: All Categories Combined CI	All Categories Combined RR (99% CL)‡	Homosexual Non-IVDAs# CI	Homosexual Non-IVDAs# RR (99% CL)	Heterosexual IVDAs¶ CI	Heterosexual IVDAs¶ RR (99% CL)
Total USA	Non-Hispanic Whites†	180,256,366	25.7	1.0	19.9	1.0	1.7	1.0
	US-Born Hispanics	9,704,157	51.7	2.0 (1.9, 2.1)	21.7	1.1 (1.0, 1.2)	20.9	12.2 (11.3, 13.1)
	Puerto Rican-Born	1,002,963	181.2	7.0 (6.6, 7.5)	45.9	2.3 (2.0, 2.6)	104.3	60.7 (55.4, 66.6)
	Mexican-Born	2,199,221	25.3	1.0 (0.9, 1.1)	18.7	0.9 (0.8, 1.1)	1.0	0.6 (0.4, 1.0)
	Cuban-Born	607,814	118.0	4.6 (4.2, 5.1)	97.2	4.9 (4.4, 5.4)	5.3	3.1 (1.9, 4.8)
	Other Latin American-Born	1,094,618	88.2	3.4 (3.2, 3.7)	64.9	3.3 (3.0, 3.6)	5.2	3.0 (2.1, 4.3)
	All Hispanics	14,608,673	70.8	2.8 (2.7, 2.8)	32.1	1.6 (1.6, 1.7)	25.3	14.7 (13.8, 15.7)
Northeast	Non-Hispanic Whites	40,995,586	31.9	1.0	22.6	1.0	5.1	1.0
	US-Born Hispanics	1,176,477	260.2	8.2 (7.7, 8.6)	60.6	2.7 (2.4, 3.0)	153.9	30.3 (27.9, 32.9)
	Puerto Rican-Born	746,193	204.0	6.4 (6.0, 6.9)	40.4	2.0 (1.7 2.3)	125.0	24.6 (22.2, 27.2)
	Mexican-Born	18,126	292.4	9.2 (6.4, 13.1)	220.7	9.8 (6.5, 14.7)	27.6	5.4 (1.7, 17.2)
	Cuban-Born	138,235	149.7	4.7 (3.9, 5.6)	125.1	5.5 (4.5, 6.7)	10.9	2.1 (1.1, 4.2)
	Other Latin American-Born	525,258	96.5	3.0 (2.7, 3.4)	67.4	3.0 (2.6, 3.4)	9.9	1.9 (1.4, 2.8)
	All Hispanics	2,604,289	241.6	7.6 (7.3, 7.9)	69.6	3.1 (2.9, 3.3)	126.8	25.0 (23.2, 26.8)
Midwest	Non-Hispanic Whites	51,510,114	8.9	1.0	6.9	1.0	0.3	1.0
	US-Born Hispanics	868,013	14.3	1.6 (1.3, 2.0)	7.3	1.1 (0.8, 1.5)	4.0	12.5 (7.8, 20.2)
	Puerto Rican-Born	103,053	87.3	9.9 (7.5, 13.0)	31.1	4.5 (2.9, 7.1)	45.6	141.5 (92.3, 216.5)
	Mexican-Born	215,247	25.6	2.9 (2.0, 4.1)	18.6	2.7 (1.8, 4.1)	0.5	1.4 (0.1, 19.1)
	Cuban-Born	21,995	72.7	8.2 (4.3, 15.6)	63.7	9.2 (4.6, 18.4)	0.0	0.0 (0.0, 79.0)
	Other Latin American-Born	68,237	33.7	3.8 (2.2, 6.5)	26.4	3.8 (2.1, 7.0)	1.5	4.5 (0.3, 60.1)
	All Hispanics	1,276,545	26.7	3.0 (2.6, 3.5)	14.6	2.1 (1.8, 2.6)	7.2	22.4 (16.0, 31.2)
South	Non-Hispanic Whites	56,028,373	22.8	1.0	17.7	1.0	0.9	1.0
	US-Born Hispanics	3,233,698	19.3	0.8 (0.8, 0.9)	13.5	0.8 (0.7, 0.9)	2.4	2.7 (2.0, 3.7)
	Puerto Rican-Born	104,422	116.8	5.1 (4.1, 6.5)	51.7	2.9 (2.1, 4.1)	37.3	43.0 (28.0, 65.9)
	Mexican-Born	527,795	11.9	0.5 (0.4, 0.7)	8.7	0.5 (0.3, 0.7)	0.6	0.7 (0.1, 2.9)
	Cuban-Born	395,351	105.7	4.6 (4.1, 5.3)	86.6	4.9 (4.3, 5.6)	3.3	3.8 (1.8, 7.8)
	Other Latin American-Born	212,700	89.6	3.9 (3.3, 4.6)	66.8	3.8 (3.0, 4.7)	0.5	0.5 (0.0, 7.1)
	All Hispanics	4,473,966	34.2	1.5 (1.4, 1.6)	24.3	1.4 (1.3, 1.5)	3.1	3.6 (2.8, 4.6)
West	Non-Hispanic Whites	31,722,293	50.3	1.0	41.2	1.0	1.1	1.0
	US-Born Hispanics	4,425,969	27.2	0.5 (0.5, 0.6)	20.1	0.5 (0.4, 0.5)	2.3	2.1 (1.6, 2.8)
	Puerto Rican-Born	49,195	168.7	3.4 (2.5, 4.4)	87.4	2.1 (1.4, 3.1)	54.9	48.4 (28.9, 80.8)
	Mexican-Born	1,438,053	26.8	0.5 (0.5, 0.6)	19.9	0.5 (0.4, 0.6)	1.0	0.9 (0.4, 1.7)
	Cuban-Born	52,233	145.5	2.9 (2.2, 3.9)	116.8	2.8 (2.0, 3.9)	7.7	6.7 (1.8, 24.6)
	Other Latin American-Born	288,423	84.6	1.7 (1.4, 2.0)	68.0	1.7 (1.4, 2.0)	1.0	0.9 (0.2, 4.1)
	All Hispanics	6,253,873	34.8	0.7 (0.6, 0.7)	25.6	0.6 (0.6, 0.7)	2.5	2.2 (1.7, 2.8)

*AIDS cases reported between June 1, 1981 and December 12, 1988, per 100,000 population of the same ethnic group (1980 census).
†Reference group is US non-Hispanic Whites (regardless of birthplace).
§Race/ethnicity and/or country of birth.
#Homosexual/bisexual men without a history of intravenous drug abuse.
¶Heterosexual men and women with a history of intravenous drug abuse.
‡CL = Confidence limits.

that were several times greater than among the other foreign-born. In the Northeast, where 64 percent of US-born Hispanics and 57 percent of all Hispanics were of Puerto Rican ethnicity,[7] most AIDS cases in US-born Hispanics and in all Hispanics were in heterosexual IVDAs. Among residents of Puerto Rico (excluded from the above analysis), the pattern was similar: 54 percent of cases were in heterosexual IVDAs.

Discussion

Although the risk of AIDS is elevated in US Hispanics,[1,2] the present study shows that the risk in persons of Mexican ethnicity is similar overall to that in the reference group of Whites who are not Hispanic, but varies by region. This finding agrees with HIV-antibody test data on applicants for military service which show that in the West (where 76 percent of Hispanics are of Mexican ethnicity), the prevalence of HIV antibody in Hispanics (0.9 percent) is the same as that in Whites who are not Hispanic (0.8 percent) (RR = 1.1, 99% confidence limits = 0.7, 1.7).**

The high proportion of cases in heterosexual IVDAs found among Hispanics overall[5] primarily reflects the high proportion in persons of Puerto Rican ethnicity (birth or ancestry). The strong association between persons of Puerto Rican ethnicity and heterosexual IVDAs with AIDS may be related to the concentration of both in the Northeast region. Of all US residents of Puerto Rican ethnicity, 74 percent resided in the Northeast,[7] where 74 percent of heterosexual IVDAs with AIDS and 89 percent of Hispanic heterosexual IVDAs with AIDS also resided. The Puerto Ricans' geographic location at the center of the AIDS epidemic in heterosexual IVDAs could have exposed them to a greater

**Unpublished data on tests done from October 1985 through March 1988, by US Department of Defense.

TABLE 2—Percentage Distribution of AIDS Cases in Hispanics and/or Latin American-Born Persons, by Exposure Category for Different Ethnic Groups and US Regions

Region	Ethnic Group*	AIDS Cases	Exposure Category of AIDS Cases: Heterosexual IVDAs†	Homosexual non-IVDAs	Homosexual IVDAs	All Others	Total
Total USA	Non-Hispanic Whites§	46,373	6.7	77.2	7.4	8.7	100.0
	US-Born Hispanics	5,015	40.4	41.9	6.3	11.4	100.0
	Puerto Rican-Born	1,817	57.6	25.3	5.6	11.5	100.0
	Mexican-Born	556	4.1	74.1	3.6	18.2	100.0
	Cuban-Born	717	4.5	82.4	4.2	8.9	100.0
	Other Latin American-Born	965	5.9	73.6	4.7	15.8	100.0
	All Hispanics	10,336	35.7	45.3	5.8	13.2	100.0
Northeast	Non-Hispanic Whites	13,075	15.9	70.8	4.3	9.0	100.0
	US-Born Hispanics	3,061	59.2	23.3	5.1	12.4	100.0
	Puerto Rican-Born	1,522	61.3	21.8	5.3	11.6	100.0
	Mexican-Born	53	9.4	75.5	1.9	13.2	100.0
	Cuban-Born	207	7.2	83.6	3.9	5.3	100.0
	Other Latin American-Born	507	10.3	69.8	3.6	16.3	100.0
	All Hispanics	6,291	52.5	28.8	5.0	13.7	100.0
Midwest	Non-Hispanic Whites§	4,564	3.6	77.8	6.2	12.4	100.0
	US-Born Hispanics	124	28.2	50.8	0.8	20.2	100.0
	Puerto Rican-Born	90	52.2	35.6	1.1	11.1	100.0
	Mexican-Born	55	1.8	72.7	1.8	23.7	100.0
	Cuban-Born	16	0.0	87.5	6.2	6.3	100.0
	Other Latin American-Born	23	4.4	78.3	8.7	8.6	100.0
	All Hispanics	341	27.0	54.8	2.4	15.8	100.0
South	Non-Hispanic Whites	12,763	3.8	77.8	7.9	10.5	100.0
	US-Born Hispanics	624	12.2	70.2	7.0	10.6	100.0
	Puerto Rican-Born	122	32.0	44.3	10.7	13.0	100.0
	Mexican-Born	63	4.8	73.0	4.8	17.4	100.0
	Cuban-Born	478	3.1	82.1	3.1	11.7	100.0
	Other Latin American-Born	191	0.5	74.4	5.2	19.9	100.0
	All Hispanics	1,528	9.1	71.1	6.0	13.9	100.0
West	Non-Hispanic Whites	15,971	2.2	81.7	9.9	6.2	100.0
	US-Born Hispanics	1,206	8.6	73.7	9.4	8.3	100.0
	Puerto Rican-Born	83	32.5	51.8	8.4	7.3	100.0
	Mexican-Born	385	3.6	74.3	3.9	18.2	100.0
	Cuban-Born	76	5.3	80.3	10.5	3.9	100.0
	Other Latin American-Born	244	1.2	80.3	6.2	12.3	100.0
	All Hispanics	2,176	7.3	73.5	8.3	10.9	100.0

*Race/ethnicity and/or country of birth.
†Intravenous drug abusers.
§For comparison with Hispanics: non-Hispanic Whites (regardless of birthplace).

risk of IVDA-related AIDS. Nevertheless, even in the regions where Puerto Ricans are a minority of the Hispanic population, the cumulative incidence of AIDS in heterosexual IVDAs was higher in persons born in Puerto Rico than in those born in Cuba, Mexico, or other Latin American countries. This suggests that the higher cumulative incidence in persons of Puerto Rican ethnicity may be due to a higher prevalence of intravenous drug abuse or sharing of needles and syringes, independent of geographic region.

Alternatively, Puerto Rican IVDAs outside the Northeast are likely to visit or be visited by other Puerto Ricans from the Northeast or from Puerto Rico itself, where most Hispanic persons with AIDS are IVDAs and the prevalence of HIV infection is high. HIV-antibody test data on applicants for military service show that the prevalence of HIV antibody in residents of Puerto Rico (0.6 percent) is almost as high as that in Hispanic residents of New York City (0.7 percent) and greater than that in residents of the United States as a whole (0.1 percent) or any of the 50 states.[8]*** Even without a higher prevalence of intravenous drug abuse, sharing needles and syringes with persons from the Northeast or Puerto Rico could account for Puerto Ricans in the Midwest, South, and West having higher cumulative incidence and higher proportions of AIDS cases in heterosexual IVDAs than other Hispanics outside the Northeast region.

Our estimates of cumulative incidence in persons born in Mexico, Cuba, and other Latin American countries may be too high, because as many as one million undocumented Hispanic aliens may not have been counted in the 1980 census.† Despite this possible bias, Mexican-born persons in the South and West had a lower cumulative incidence than White non-Hispanics. This possible bias is less likely to apply to persons born in Puerto Rico, as their US citizenship gave them no motive to avoid participation in the US census.

We have no data on the country of ancestry of US-born Hispanics with AIDS. It would be useful to know whether patterns of AIDS risk according to ancestry follow those according to birthplace. The high cumulative incidence and high proportion of cases in heterosexual IVDAs among

***Unpublished data on tests done from October 1985 through December 1987, US Department of Defense.

†Personal communication: Karen Woodrow, Bureau of the Census.

US-born Hispanics in the Northeast (where 57 percent of the Hispanic population is Puerto Rican) and the low cumulative incidence and low proportion in heterosexual IVDAs among US-born Hispanics in the West (where 76 percent of the Hispanic population is Mexican) suggest that the patterns are similar. If so, strategies for preventing AIDS[9–11] in US-born Hispanics should be similar to those for non-US-born Hispanics. Our data for the latter suggest that resources for preventing AIDS in Hispanics are needed most in those of Puerto Rican ethnicity, particularly for AIDS related to intravenous drug abuse.

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3rd International Conference on AIDS Education: Call for Abstracts Extended to July 7

"Changing Environments: Meeting the Challenge" is the theme for the 3rd International Conference on AIDS Education, to be held September 10–13, 1989 in Nashville, Tennessee. The conference is sponsored by the International Society for AIDS Education and several other international and national organizations.

The main purpose of the conference is to bring together persons from academic, governmental, and community settings who are working in the interdisciplinary field of AIDS education, counseling, and prevention. Conference participants will review current approaches and methods in the context of meeting the challenges posed by changing social, economic, and institutional environments, as well as changing communication and information technologies. The adaptation of AIDS education and counseling programs to such change will be emphasized. Policy, psychosocial, legal, and ethical issues will be addressed.

The format of the conference will provide structured and unstructured periods for networking and idea exchange as well as plenary sessions featuring experts in the field. Concurrent sessions will be classified into three groups: Research Seminars, Roundtable Discussions, and Workshops. A Poster Session will further enhance knowledge and networking. Informal networking will occur during breaks, social events, and in the exhibits area.

The call for abstracts has been extended to include abstracts postmarked by July 7, 1989. To obtain an abstract form or additional information, contact: Conference Secretariat, 3rd International Conference on AIDS Education, P.O. Box 582, Vanderbilt University Medical School, Nashville, TN 37232-2351 USA. Tel: (615) 322-2437 or (615) 433-2252.

HIV Testing in Prisoners: Is Mandatory Testing Mandatory?

Jon K. Andrus, MD, David W. Fleming, MD, Catherine Knox, MN, Robert O. McAlister, PhD, Michael R. Skeels, PhD, MPH, Robert E. Conrad, PhD, John M. Horan, MD, MPH, and Laurence R. Foster, MD, MPH

Abstract: We studied 977 newly incarcerated Oregon inmates to compare voluntary versus mandatory human immunodeficiency virus antibody (HIVAb) testing in the prison setting. All inmates were offered HIVAb counseling and testing. Blood drawn for routine syphilis serology from those who declined this offer was also tested for HIVAb after personal identifiers had been removed. Only 1.2 percent (12) prisoners were HIV positive. However, 62.5 percent (611) inmates were at risk for HIV infection by being an intravenous drug user, a male homosexual, or hepatitis B core antibody (HBcAb) positive. The ratio of at-risk, as yet uninfected inmates to those already HIV infected was 53 to 1. Two-thirds of all inmates including those at-risk chose to receive counseling and testing. In areas where most at-risk inmates are not yet infected, it may be more appropriate for HIV prevention activities in prison to focus on voluntary programs that emphasize education and counseling rather than mandatory programs that emphasize testing. (*Am J Public Health* 1989; 79:840–842.)

Introduction

Persons for whom HIVAb (human immunodeficiency virus antibody) testing is already mandatory in the United States include immigrants, blood donors, and military recruits. Prison inmates comprise a population at risk for HIV infection. Many are intravenous (IV) drug users and all are entering a setting where homosexual sex may occur. The American Medical Association has recommended that "testing for AIDS virus should be mandatory . . . for inmates in federal and state prisons."[1] Currently, inmates in all federal prisons are tested. In 1986, three state prison systems required their prisoners to be tested for HIVAb; by 1988 this number had increased to 14.[2,3] However, little data have been available to evaluate the usefulness of mandatory HIVAb testing in the prison setting.

In 1987, the Oregon Health Division performed a study to evaluate the usefulness of a voluntary versus a mandatory HIVAb testing strategy in the Oregon corrections system.

Methods

All newly incarcerated inmates entering any of the five prisons of the Oregon State correctional system from September 1987 to January 1988 were enrolled in the study. Upon entry, each inmate was shown a 12-minute educational videotape on HIV infection and was then administered a standard questionnaire in private by a Corrections Department health care provider. Information collected from each inmate included: history of blood transfusion, sexual preference, IV drug history, self-assessed likelihood of HIV exposure, and HIV testing history. Additional information extracted from the records of inmates included: age, gender, race, education, crime, length of sentence, and history of sexually transmitted diseases.

All inmates were then offered confidential HIV antibody testing and counseling. Inmates were categorized into those choosing testing and those declining confidential testing. Inmates who declined were asked whether they would consider testing in the future if it could be offered in such a way that only they had knowledge of the results (anonymously). Those who would have considered anonymous testing were grouped separately from the rest of test decliners in some of the analyses.

Blood specimens from inmates choosing testing were tested for HIV antibody and results were presented to the inmate, along with post-test counseling. Blood specimens from all inmates are routinely drawn and tested for evidence of syphilis infection. Using these sera, a 95 percent random sample of all inmates declining testing was also tested for HIVAb in a blinded fashion after syphilis testing had been completed. These HIVAb results, with linked questionnaire data, were sent only to the Oregon Health Division. The study design was approved by the Oregon Health Division's human subjects review board, which included a medical ethicist, a lawyer from the American Civil Liberties Union, and representatives from prisoners, gays, and a community AIDS-prevention agency.

All specimens were also tested for hepatitis B core antibody (HBcAb). Hepatitis B and HIV infections are transmitted in the same way; we used HBcAb positivity as an objective, surrogate marker for a history of risk behavior for HIV infection. For purposes of this study, inmates were considered to be at risk for HIV infection if they had one or more of the following: IV drug use history, history of male homosexuality, or HBcAb positivity.

Serum specimens were tested at the Oregon Public Health Laboratory. HIV antibody testing was done using the Genetics Systems* enzyme linked immunosorbent assay. Repeatedly reactive sera were confirmed by the indirect immunofluorescence test. HBcAb tests were done using the Abbott* enzyme linked immunosorbent assay.

Black, Hispanic, and Native American inmates were grouped together as minority inmates. Male inmates were defined to have admitted homosexuality if they stated that they had ever had sex with men only or with both men and women.

All statistical tests were performed using the Epi Info computer software package provided by the Centers for Disease Control.

From the Centers for Disease Control, Division of Field Services, Epidemiology Program Office (Andrus, Horan); Oregon Health Division, Portland (Fleming, McAlister, Skeels, Conrad, Foster); and Oregon Department of Corrections, Salem (Knox). Address reprint requests to Jon K. Andrus, MD, Division of Field Services, Epidemiology Program Office, Centers for Disease Control, 1600 Clifton Road, Atlanta, Georgia 30333. This paper, submitted to the Journal December 5, 1988, was revised and accepted for publication March 21, 1989.

*Use of trade names does not imply endorsement by the Oregon Health Division or by the US Public Health Service.

Results

Seroprevalence

A total of 995 consecutively admitted inmates were enrolled; 18 of the test decliners were randomly excluded from HIVAb testing. The mean age of the remaining 977 was 30 years (range 17 to 71 years), 91 percent male, 74 percent were White, 16 percent Black, 6 percent Hispanic, and 2 percent Native American.

Twelve inmates (1.2 percent) were HIVAb positive (Table 1).

Inmates who were HBcAb positive were more likely to be HIV infected (2.5 percent, 9/357) than those who were seronegative for HBcAb (0.5%, 3/620) (RR = 5.2; 95% CL 1.6, 16.7). Persons who admitted homosexuality or IV drug use were not more likely to be HIVAb positive than those who did not admit these behaviors, although the small number of HIV-seropositive inmates limited our power to detect a difference. The risk of HIV infection was not associated with other inmate characteristics studied, including race, education, history of previous HIV testing, crime, length of sentence, history of sexually transmitted diseases.

Risk Behavior and HIV Seroprevalence

Of the 977 inmates enrolled, 611 (63 percent) were HBcAb positive or reported a history of IV drug use or reported history of male homosexuality, and thus were at risk for HIV infection. This 63 percent does not include inmates who reported only a history of blood transfusion or sexually transmitted disease. Thus, 63 percent may be a conservative estimate of the proportion of inmates at risk for HIV infection. All 12 HIV positive inmates were in this group defined to be at risk. For every inmate infected, there were 53 others at risk, but as yet uninfected.

Test Choice

A total of 637 inmates (65 percent) chose to receive HIV counseling and testing. Of the 358 inmates declining testing, 98 (27 percent) would have considered anonymous testing, if it had been available.

Six of the 12 HIV-positive inmates voluntarily accepted an offer of confidential counseling and testing. HIV-positive inmates were not significantly different from HIV-negative inmates in the likelihood that they would choose to be tested confidentially. Among the 340 test decliners, 67 percent (4/6) of HIV-positive inmates would have considered anonymous testing compared to 28 percent (94/334) of HIV-negative inmates (RR = 2.4; CL 1.0, 5.4).

Test choice was not predicted by HBcAb status, admitted male homosexuality, age, gender, race, education, crime, and sentence. Seventy percent (361/519) of inmates who admitted to IV drug use accepted an offer of testing compared to 60 percent (276/457) of inmates who denied IV drug use (RR = 1.2; 95% CL 1.1, 1.3). Eighty-five percent (63/74) inmates who believed it likely that they had been exposed to HIV chose testing compared to 64 percent (566/885) of those who either did not believe or were not sure about HIV exposure (RR = 1.3; 95% CL 1.1, 1.6).

In this study, 35 percent (340/977) of the inmates declined the offer of confidential counseling and testing. Of those who declined, 56 percent (192/340) were at risk by being HBcAb positive, an IV drug user, and/or a male homosexual.

TABLE 1—HIV Seropositivity in Inmates by Demographic Characteristics and Risk Factors, Oregon, 1988

	Enrolled	HIV Seropositive (%)
Total	977	12 (1.2)
Sex		
Male	893	11 (1.2)
Female	82	1 (1.2)
Unknown	2	0 (0.0)
Age (years)	Mean 30 Range 17–71	Mean 31 Range 29–38
Race		
White	722	6 (0.8)
Black	161	3 (1.9)
Hispanic	61	2 (3.3)
Native American	23	1 (4.4)
Unknown	10	0 (0.0)
Risk Factors		
IV Drug User	519	10 (1.9)
HBcAb Positive	349	9 (2.6)
Male Homosexual	30	1 (3.3)
Unduplicated Total	611	12 (2.0)

Discussion

Both mandatory and voluntary testing strategies have their inherent limitations.[4,5] If the goal of testing is to identify all infected inmates so that they can be quarantined, then testing must be mandatory.[6] However, in Oregon, prison officials decided to use a health facility model of infection control and to implement universal HIV precautions rather than try to identify and quarantine infected inmates. One reason for this decision was that the average prison stay in Oregon is only 16.8 months (Oregon Department of Corrections, Statistical Report, PP30TC11, May, 1988). A policy that isolates HIV positive inmates while they are imprisoned does little to prepare either the prisoner or society for the near future when they are released.

Given these preconditions, the primary reason to test inmates in Oregon would be to inform them of the results, in hopes that this information would help them to implement risk reduction behavior. The data from this study provide insight into the kind of testing program that would most likely accomplish this goal.

The prevalence of HIV infection among prisoners entering the Oregon correctional system at the end of 1987 was 1.2 percent. While this estimate is relatively low compared to infection rates of 15 percent–20 percent reported in prison populations in high AIDS incidence areas,[3,7,8] it is probably more representative of that expected in most states. Oregon is nineteenth among states ranked by cumulative reported AIDS incidence rate.[8] To date, prisoner seroprevalence data have been reported from 20 states; in 17 of the states (85 percent), seroprevalence was less than 2 percent.[2]

In contrast to the low HIV seroprevalence, 63 percent of inmates entering Oregon corrections system were either HBcAb positive or admitted to IV drug use or male homosexuality, all indicators of potential risk for future HIV infection. The ratio of uninfected to infected at-risk inmates was 53:1.

The problem then is not that there are a lot of infected inmates that need to be identified. Rather, there are a lot of at-risk, as yet uninfected inmates that need to be convinced to change their behavior. To be effective, prevention programs must be specifically directed at this latter group.

To our knowledge, the influence of mandatory HIV testing on subsequent behavior has not been evaluated. It seems particularly unlikely, however, that forced testing would significantly change the conduct of those testing negative. In

contrast, multiple studies have documented that voluntary HIV testing and counseling can change behavior.[9–15] More importantly, this intervention has specifically been shown to cause the at-risk client who tests negative to initiate risk reduction behavior.[12–14] When most persons at risk are not yet infected, voluntary HIV prevention programs that emphasize counseling may be more effective than mandatory programs that emphasize testing.

Our study demonstrates that such a voluntary program can be successfully implemented in the prison setting. In Oregon, two-thirds of all inmates, including those at highest risk for HIV, sought HIV counseling and testing when given the opportunity.

In conclusion, we believe this study shows that an effective HIV intervention program for Oregon's prisoners should contain the following components: general AIDS education for all prisoners to raise their basic awareness of the disease; counseling at the prisoners request, especially for those who indicate they may be at risk; and HIV testing, offered only as voluntary component of this counseling, so that testing cannot occur without counseling, but counseling can occur without testing. Anonymous HIV testing was attractive to some inmates who declined confidential testing and when feasible should be considered as a testing option.

The Oregon Legislature has accepted these recommendations and approved the funding necessary to implement a comprehensive voluntary HIV counseling and testing program in Oregon prisons.

ACKNOWLEDGMENTS

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Task Force Report on Clinical Preventive Services

The US Preventive Services Task Force recently completed a five-year review of numerous clinical interventions and presented their recommendations to Dr. J. Michael McGinnis, deputy assistant secretary for health (disease prevention and health promotion). Their report, entitled *Guide to Clinical Preventive Services*, was unveiled May 2 during a news conference in Washington, DC.

The report is the work of a 20-member non-federal panel appointed in 1984 to develop age- and risk factor-specific recommendations for the delivery of preventive services in the clinical setting. The report addressed the effectiveness of over 100 clinical interventions for the prevention of 60 diseases. Its basic recommendation is that clinicians should tailor preventive examinations to the individual patient rather than routinely performing a battery of tests on patients which may be ineffective and could lead to inaccurate results.

The report recommends four types of interventions in the clinical setting: screenings, counseling, immunizations, and chemoprophylaxis. Of these, said Robert S. Lawrence, MD, who chaired the task force, counseling patients about behavior modification can be among the most effective forms of preventive, adding that ". . . the key to improved health and disease prevention often lies in lifestyle changes made by patients rather than in the broad testing or sophisticated medical procedures performed by health providers."

The all-volunteer task force was comprised of 14 physicians as well as specialists in dentistry, nursing, health services research, health education, health economics, and medical sociology. The report—Guide to Clinical Preventive Services—will be published by Williams and Wilkins of Baltimore, MD in late summer. For further information, contact the publisher at 1-800-638-0672.

Emergency Room Use and Primary Care Case Management: Evidence from Four Medicaid Demonstration Programs

ROBERT E. HURLEY, PHD, DEBORAH A. FREUND, PHD, AND CAPT. DONALD E. TAYLOR, USAF

Abstract: Claims-based utilization data collected in four Medicaid demonstration programs are used to examine the impact of primary care case management on patterns of reliance on the emergency room as a source of services. The experience of stratified random samples of AFDC (Aid to Families with Dependent Children) adults and children in the demonstration programs is compared with that of equivalent samples from comparison groups in traditional Medicaid programs. Data were analyzed in multivariate models controlling for person characteristics in a pre/post, demonstration/comparison design. Results indicated large reductions in the proportion of persons with at least one emergency room visit ranging from 27 to 37 per cent for children and 30 to 45 per cent for adults. Use levels for persons with at least one ER visit are less substantially affected. The findings suggest that gatekeeping designs can alter patterns of enrollee use with respect to the emergency room and some discussion of these implications is presented. (*Am J Public Health* 1989; 79:843–847.)

Introduction

Access to health care by low income Americans has been of considerable policy concern for the last two decades. Of special import has been the finding that poor individuals are less likely than persons of more substantial means to have a regular source of care. Thus, for a variety of institutional and economic reasons, these persons are more likely to use emergency rooms instead of private physicians for their basic medical care.

Primary care case management programs offer promise for reducing reliance on the emergency room as a regular source of primary care.[1] By channeling access to all medical services through a primary care provider, these programs make available other sources of care and provide constraints on the free access to the emergency room for non-emergency services. This two-pronged approach has been of special interest to state Medicaid agencies concerned about improving access to care and using program dollars more cost efficiently.[2,3]

A substantial base of research has been accumulated to demonstrate that the majority of services rendered in emergency rooms has been for non-urgent care[4–7] and that low income persons are disproportionate consumers of primary care in the emergency room.[8] These patterns of use have been attributed to the accessibility of the emergency room during non-traditional hours of operation; its relative proximity to care-seekers brought on by the maldistribution of primary care providers; or absence of alternative sources due to ability to pay.[8] For state Medicaid agencies this reliance on the emergency room has been shown to lead to higher costs.[9]

Complementing these findings is a large literature indicating that the availability of a regular source of care from alternative sites such as office-based physicians or health centers can be effective in diverting care from the less desirable emergency room.[10,11] In an era characterized by severe retrenchment in Medicaid programs, reductions in emergency room use might be expected to produce real savings without reductions in access, if the availability of alternative sources of care can be ensured.

In this study, we report results from four Medicaid demonstration programs that, among other goals, targeted emergency room use for reductions via primary care case management.

Methods

Study Settings

The Health Care Financing Administration (HCFA) approved eight demonstration programs in six states in 1982 that are referred to as the Nationwide Medicaid Competition Demonstrations.[12] These demonstration programs incorporated components of capitation, case management, and limitation on freedom of choice.[2] Four of these programs have been examined extensively in the HCFA-funded evaluation contract awarded in 1983.[13] The focus of the evaluation was on the impact of the programs on the utilization and cost experience of enrolled beneficiaries in each demonstration program in comparison with the experience of equivalent samples of non-enrollees.

We studied two county-based demonstration programs in California and state operated capitation programs in Kansas City, Missouri and the state of New Jersey.[12] All four of these programs used a gatekeeper (case manager) model in which eligible persons were enrolled with a participating individual physician or primary care organization. The gatekeeper was required to provide all primary care and to authorize referral care, including all inpatient and non-urgent emergency room use. Initially enrollees were permitted to select their case manager in each site but were limited to using their selected provider thereafter.

In all four programs, the state Medicaid agency paid the entities with which it contracted on a prepaid basis. The capitated rate differed for each public assistance aid category included in the demonstration. In California, the contractors were county-based, quasi-governmental authorities which in turn negotiated contracts with medical care providers within the demonstration county to provide primary care case management. The two California and the Missouri programs were mandatory for Medicaid eligible persons in the covered categories while the New Jersey demonstration was a voluntary enrollment program.

Several other differences existed among the four programs. The Santa Barbara (CA) program paid case managers on a capitated basis for primary care services while the

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Monterey (CA) program continued to pay them on a fee-for-service basis plus a nominal monthly case manager's fee. Both of the California programs permitted some institutional providers to participate, i.e., health centers and hospital clinics, though the preponderance of case managers were individual physicians. Primary care providers in both California demonstrations were required to share financial risk with the intermediary organization, by participating in savings and losses resulting from specialist and emergency room use that was different from predicted levels. Case managers in New Jersey were paid on a capitated basis for primary care services similar to Santa Barbara.

The Missouri Medicaid program engaged five organizations and paid them on a capitated basis for covered services: two neighborhood health centers, two hospital outpatient clinics, and an IPA-model health maintenance organization.[14] Only in the IPA were enrollees assigned to a specific primary care case manager while in the remaining providers the organization was, in effect, the case manager. Missouri also permitted approximately 50 individual physicians to perform case management for their existing Medicaid patients. These physicians continued to be paid on a fee-for-service basis with a supplemental case management fee paid each month for each enrollee, similar to the arrangement in Monterey county in California.

Given the basic similarities in program designs with respect to prior authorizing non-urgent emergency room use, all of the programs were expected to effect reductions in levels of emergency room use by lowering the number of unnecessary visits in the emergency room and by increasing access to alternative sources of primary care. We hypothesized these reductions would be detected for both adult and children enrollees. Furthermore, reductions were expected to be detectable in both the probability that an enrollee would have at least one visit and in the mean number of visits for enrollees with at least one visit.[15]

Data Source

Person-level analysis files have been constructed from Medicaid eligibility and claims records for Medicaid beneficiaries residing in the demonstration counties and for equivalent samples residing in selected comparison counties in California and Missouri. Ventura County was chosen because of its similarities with and geographic proximity to both California demonstration counties. St. Louis was selected because it is the only other major metropolitan county in Missouri. Since enrollment in New Jersey was statewide and voluntary, we selected a comparison group of non-enrollees from persons who were never enrolled in the demonstration program. We drew random samples of approximately 3,000 adults and children stratified by eligibility category and age from eligibility files for the year prior to implementation and for the first full year of the demonstration programs. Cost and analytical considerations precluded us from selecting a panel design and we have used a pre/post non-equivalent comparison group design to analyze our two cross-sections of data.

In the pre- and post-periods for both sites, we used claims data collected in the Medicaid Management Information Systems of each state to construct data on emergency room use. In Missouri, the claims data were partially comprised of "pseudo-claims" submitted by the capitated plans who were responsible for paying emergency rooms directly for visits by their enrollees and then submitting a pseudo or no-pay claim to the Medicaid agency. In all other sites, emergency room visits were paid directly by the state Medicaid agency.

We developed an algorithm to transform claims files into visit (event) files.* An emergency room visit is defined as a medical event that occurred in the emergency department of a hospital. Medical events were created by accumulating all medical services provided to the same person on the same day into a single measure of use. Place of service is determined by the location codes on the claims received for the services accumulated into a particular event. An indicator of having had at least one emergency room visit and the count of visits for persons with at least one such visit were accumulated to the person level records built from the samples described above. We present results for the AFDC population only since this was the only eligible group for whom results were available from all four sites.

Analysis

For each person in each of the four demonstration-comparison contrasts, we collected information on age, sex, race (not available in California), total months of eligibility in the studied year, whether a person had insurance coverage beyond Medicaid (not available in Missouri), and whether the person had continuous eligibility through the study year; in addition to these personal characteristics, variables included presence in the demonstration or comparison group, and in the pre-year or demonstration year samples. Finally, a program effect variable was modeled as an interaction term of pre/demonstration year and demonstration/comparison county and coded as follows: 1 if in the demonstration county in the demonstration year and 0 otherwise. The multivariate approach is described in greater detail in the Appendix. Separate analyses were conducted for children and for adults.

In each of the four contrasting sites, we developed estimates of the effects of each of these independent variables on the probability of use and the volume of use for users employing logit and ordinary least squares regression, respectively. While probit is the preferred statistical application with probability of use,[15] we selected the less costly and more easily retransformed logit approach since it produces virtually identical results in this case.[16]

Results

Table 1 presents our estimates of the regression-adjusted proportion with at least one visit and the percentage change in this proportion from the pre- to the demonstration year for both the demonstration group enrollees and the comparison group enrollees for AFDC mothers and children separately.

There is close comparability between demonstration and comparison groups across all four contrasts in the year prior to the demonstrations with emergency room visits higher in the two urban Missouri counties and the New Jersey enrollee and non-enrolled groups than in the less urban California counties.

The change in proportion of Medicaid beneficiaries with a visit is relatively small from the pre-year to the first year of the demonstration for comparison group members in each contrast. However, substantial reductions are noted in all of the demonstration programs in the demonstration year. For children, these reductions range from 27.5 per cent in Santa Barbara to 36.7 per cent in New Jersey. For adults the reductions were even larger ranging from a low of 30.6 per

*Hurley R, Frick G: "E Pluribus Unum: Designing Uniform Analysis Files from Multi-Year, Multi-State Medicaid Data Files." Paper presented at the Annual Meeting of the American Public Health Association in New Orleans, LA, October 1987.

TABLE 1—Regression Adjusted Estimates of Emergency Room Use (Demonstration site versus Comparison site)

Demonstration vs Comparison Sites	Proportion with at least One ER Visit					
	Children			Adults		
	Pre[A]	Demo	Per cent Change	Pre	Demo	Per cent Change
Monterey, CA	18.0	11.6	−35.6	22.8	14.0	−38.6
Ventura, CA	20.5	21.7	+ 5.9	21.7	23.1	+ 6.5
n	7,685			4,738		
Santa Barbara, CA	19.3	14.0	−27.5	23.2	14.1	−30.6
Ventura, CA	19.2	19.2	—	25.8	27.7	+ 7.4
n	7,500			4,923		
Kansas City, MO	40.7	26.5	−34.9	37.7	21.1	−44.1
St. Louis, MO	37.2	37.0	—	39.2	34.0	−13.3
n	11,031			5,001		
NJ Enrollees	37.7	23.9	−36.7	32.1	18.2	−44.3
NJ Non-enrollees	32.0	35.2	+ 9.7	31.0	33.4	+ 7.7
n	8,444			3,999		

	Monterey	Santa Barbara	Kansas City	New Jersey
[A] Pre-year	CY 82	CY 82	CY 83	FY 84
Demo-year	FY 84	FY 84	FY 85	CY 85

cent in Santa Barbara to 44 per cent in both Kansas City and New Jersey.

The analysis of program effects for persons with at least one visit found much weaker evidence of gatekeeping impact on emergency room users. Significant reductions were found for AFDC children in Monterey, Santa Barbara, and New Jersey but not in Missouri. Although reductions were detected for adult users in all of the demonstration sites, none of the interaction terms produced a p value less than .05. Because emergency room use is a relatively infrequent event, significant changes for users may be more difficult to detect with correspondingly smaller sample sizes.

Discussion

These results suggest the gatekeeping effects of primary care case management programs are strong and pervasive with respect to likelihood of an emergency room visit. Despite substantial variation in program designs, the effects were noted in each demonstration program for both children and adults. The findings should, however, be tempered by the fact that our cross-sectional design is unable to control for all potentially confounding factors that might influence the results.

While these findings confirm the expectations of the state and federal officials responsible for designing these demonstrations, the magnitudes of the effects are surprising for three reasons:

- Despite the gatekeeping design, case managers had broad discretion to authorize emergency room use and could choose to exercise this discretion either loosely or rigorously.
- Emergency rooms in some sites expressed great reluctance to deny care to on-site patients for whom authorization could not be obtained,[12] indicating direct gatekeeping effects would be attenuated in these sites.
- The results presented from each site are from the first year of demonstration program operation indicating that learning effects associated with greater comprehension of the implication of gatekeeping would have been limited during this period.

The sharp reductions in probability of a visit for adults and for children in the first year suggest that the gatekeeper concept was apparently reasonably well understood by enrollees. They may also suggest that the requirement of around-the-clock availability of the primary care case manager provided an accessible alternative to the emergency room. The data presented here do not permit us to attribute the reduction specifically to either or both of these program features and this issue is still under study in the demonstration evaluation. The fact that use for users was not so dramatically affected as probability of use suggests gatekeeping for individuals with a propensity to use the emergency room is less effective or perhaps less suitable.

Another interesting implication of the results presented here is the effectiveness of primary care case management in reducing emergency room use irrespective of the method by which the case manager is paid. For example, the fee-for-service Monterey program had gatekeeping effects similar to those found in Santa Barbara and New Jersey, which capitated primary care, and in Kansas City, which capitated nearly all covered services. This finding is especially notable given the notoriety achieved by Monterey with its subsequent dissolution in insolvency and its widespread characterization as a demonstration program failure.[17] There is also no apparent difference in effects on probability of a visit and use for users that is attributable to the method of case manager payment. Capitation does not appear to be any more effective in reducing use for users than does fee-for-service, as might be hypothesized.

The data presented here are silent on a critical question related to emergency room gatekeeping: is it non-urgent emergency room use that is being reduced in the demonstration programs? Because the program designs do not require prior authorization for bona fide emergencies, we may suspect that emergency room reductions would not be or, at least, need not be occurring for such conditions. Furthermore, the National Medical Care Utilization and Expenditure Survey (NMCUES) estimated that 85 per cent of emergency room visits are for non-life threatening reasons.[18] Thus, despite the large reductions attained in emergency room use in the demonstration programs, we may infer that levels of use in the programs still include a substantial amount of primary care.

In a detailed study of the Kansas City demonstration using data from a consumer survey,[19] we have attempted to

assess if remaining emergency care was "necessary" using both patient-assessed and clinical-based measures. We concluded that lower levels of use were associated with a higher proportion of remaining visits being clinically necessary. Unfortunately, the claims-based data employed here do not permit this analysis to achieve the same type of detail. However, other studies from related data bases indicate that access to and satisfaction with care for enrollees and nonenrollees were comparable in Missouri and California.[20] A quality of care study using outcome data has found no evidence to suggest adverse consequences associated with reductions in emergency room visits.[21]

It is notable that while emergency room use has been reduced, primary care visits in most sites did not increase correspondingly as might have been expected if substitution of such care were occurring. This finding may indicate an improvement in continuity of care fostered by formal enrollment with a responsible case manager. Other results from the evaluation have demonstrated that there are significant increases in the concentration of enrollees' care with their case managers.[22] If access to care has not been impeded and if quality of care has not changed, then these capitated demonstration programs appear to be capable of producing meaningful improvement in the continuity of care for beneficiaries, as well as real, although modest, cost savings for Medicaid agencies.

APPENDIX

Multivariate Approach Used in the Medicaid Competition Demonstrations Evaluation

Analysis of covariance techniques were used with dummy variable specifications of the time periods and demonstration versus comparison groups. The generic specification of the regression model is as follows:

$$Y = B0 + B1(DEMCOMP) + B2(PREPOST) + B3(PROGRAM) + BkX + U$$

where:

Y = specific dependent utilization measures under examination;

DEMCOMP = a dummy variable indicating if the observation came from the demonstration or comparison group (coded 1 if demonstration; 0 if otherwise);

PREPOST = a dummy variable indicating if the observation came from the pre- or post-(demonstration) year period (coded 1 if post-year; 0 if otherwise);

PROGRAM = a term defined as DEMCOMP × PREPOST used to indicate if observation came from demonstration group in the demonstration year (coded 1 if both DEMCOMP and PREPOST = 1; 0 if otherwise);

X = a vector of control variables, including demographic and eligibility characteristics;

U = the error term.

B0 and B1 to Bk constitute the intercept term and the regression coefficients being estimated that lead to the estimate of program effects.

The parameter estimates for DEMCOMP, PREPOST, and PROGRAM are summarized in the Appendix Table.

Appendix Table

		Site		Difference
		Comparison	Demonstration	(Demo–Comp)
Time Period	Pre	B0	B0+B1	B1
	Post	B0+B2	B0+B1+B2+B3	B1+B3
Difference (Post-Pre)		B2	B2+B3	B3

To generate the regression adjusted proportions with an emergency room visit shown in Table 1, predicted values for each site and year were produced controlling at the mean levels on the included control variables. Further detail on the analytical approach including the full regression results are available from the authors.

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Determinants of Community-Based Coverage: Periodic Vitamin A Supplementation

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Abstract: Factors related to preschool child receipt of vitamin A during the first year of a semi-annual vitamin A capsule delivery program were investigated in 229 villages in Aceh, Indonesia. Coverage was higher in villages which were more rural and less economically developed. Highest performance was achieved by village distributors who represented the local status quo in this rural area (farmers, or non-farmers with minimum education) rather than more upwardly mobile, highly educated residents. Household or child-level characteristics were not associated with coverage. This information may be useful for planning direct service programs in the community. (*Am J Public Health* 1989; 79:847–849.)

Introduction

Vitamin A deficiency is the major cause of pediatric blindness in developing regions throughout the world.[1] Milder systemic vitamin A deficiency may also predispose children to greater risk of infectious morbidity[2] and mortality[3–5] and appears to be an important determinant of child growth.[6,7]

One major strategy for preventing vitamin A deficiency in the community is through periodic delivery of oral doses of vitamin A (usually 200,000 IU every four to six months) to preschool age children. Distribution can be carried out either as a special, single-purpose program or integrated into existing primary health care activities.[8] Presently, over 100 million oral doses of vitamin A are distributed annually, the bulk of which are provided as gelatin capsules by UNICEF.[9]

Community and distributor factors influencing capsule coverage are largely unknown. These characteristics were investigated during a semi-annual, large-dose vitamin A supplementation program in Indonesia from 1982–84.[10]

Methods

This evaluation formed part of a large community trial designed to assess the impact of vitamin A supplementation on childhood xerophthalmia, mortality, and growth in the north Sumatran Province of Aceh—population: 3.87 million; area: 55,390 sq. kilometers. The basic study design and findings on mortality have been reported.[4] Two-hundred-twenty-nine villages were randomly selected to receive the Indonesian Government's single-purpose vitamin A capsule distribution program operated through the provincial health service. Village distributors were identified based on village headmen recommendations and a minimum fluency in reading and writing the Indonesian language. Two days' training were given at local health centers concerning the purpose of vitamin A distribution, program organization, house-to-house visiting and oral dosing procedures, record keeping and reporting. Distributors were paid an honorarium equivalent to US $2.00 (1983) for each six-monthly distribution, per government guidelines. Normal supervision was strengthened by adding two study monitors who visited program villages and encouraged as complete coverage as possible.

Village, distributor, household, and child characteristics were assessed during a baseline survey by two field teams. Capsule distribution began within two months after the survey, and again six months later. Coverage, assessed one year after the baseline survey, was considered adequate if children received a vitamin A capsule at both capsule distribution cycles. Stratified continuous and discrete data were analyzed by the chi-squared test.[11] A multiple linear regression model was fit to data to examine the adjusted effects of selected factors on coverage.[12] Data were analyzed using SAS (SAS Institute, Cary, NC).

Results

Distributors selected by village leaders differed from heads of household in program villages (from which they were likely to have been drawn) (Table 1) with larger proportions being younger, female, and better educated (possibly due to the program's minimum literacy requirement).

More than half (54 per cent) of all program distributors achieved high, 30 per cent moderate, and 16 per cent low coverage of children at both distribution cycles (Table 2). Education was inversely related to distributor performance. Farming was related to higher coverage compared to non-farming occupations (mostly government workers, self-em-

TABLE 1—Comparisons Between Capsule Distributors and Adult Heads of Household in Program Villages, Aceh 1982–84

Characteristics	Distributors (%)	Heads of household
Number	211*	8,246
<40 years of age	71.1	56.2
Male sex	85.8	97.1
>6 years education	55.7	19.9
Farming occupation	61.7	53.9

*Excludes 18 with missing values. Distributions of education and occupation based on 210 and 193 values, respectively.
†Includes landowning and tenant farmers.

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TABLE 2—Capsule Coverage by Distributor and Village Characteristics

Characteristics	No.	%	Per cent Coverage <50%	50–79%	80–100%	P
Distributor Characteristics						
Education (years)						
≤6	93	44.3	9.7	30.1	60.2	
>6	117	55.7	20.5	29.1	50.4	<0.100
Occupation						
Farming	145	61.7	15.1	24.4	60.5	
Other	48	38.3	18.9	37.8	43.2	<0.050
Cropland owned						
None	92	43.6	10.9	32.6	56.5	
<½ ha	40	19.0	35.0	25.0	40.0	
≥½ ha	79	37.4	11.4	29.1	59.5	<0.010
Village Characteristics						
Village Size (No. houses)						
<30	79	37.6	17.7	17.7	64.6	
30–59	106	50.5	13.2	35.9	50.9	
≥60	25	11.9	20.0	20.0	40.0	>0.100
Distance to major road						
<1 km	54	23.7	27.8	42.6	29.6	
≥1 km	174	76.3	14.4	25.3	60.3	<0.001
Shop in village						
Yes	125	66.8	13.6	39.2	47.2	
No	62	33.2	17.7	16.1	66.1	<0.010
Total	229	100.0	15.6	29.9	54.5	

Note: 19, 36, 18, 19, 1, and 42 missing values for distributor education, occupation, cropland owned, village size, distance to major road, shop in village, respectively.

"Coverage" defined as the percentage of children in a village receiving a vitamin A capsule at both distribution cycles.

ployed, and craftsmen). Cropland ownership was associated in a somewhat unexpected fashion to coverage: distributors who owned only a small amount (0 ha < X <½ ha) performed poorly compared to those who owned either ≥½ ha or no cropland at all. Other distributor characteristics such as age, sex, ownership (bicycle, motorcycle, or radio), or relationship to the village headman were unrelated to distributor performance (not shown).

Capsule coverage was weakly associated with the size of a village (number of houses), but inversely related to a cluster of village "development" factors. Strongest was the association with proximity to a major (district or subdistrict) road, with twice the proportion (60 per cent) of more distant villages (≥1 km) having high coverage vs those in closer proximity (<1 km, 30 per cent). Similar associations were noted with distances to other development landmarks such as local health centers, weekly markets, and elementary schools (not shown). Lower coverage villages were more likely to have at least one commercial shop.

Multivariate adjustment by least squares linear regression confirmed the weak influence of village size, the strong positive effect of a village's distance from a major road, and the strong, negative effect of a distributor owning a small amount of cropland on coverage (Table 3). The influences of distributor occupation and education were manifest through an interaction: more education plus a non-farming occupation was associated with markedly poorer coverage than any of the other three joint categories of occupation and education. Further stratified analyses showed that 60 per cent of farmers achieved high coverage (regardless of education) as well as 55 per cent of poorly educated (<6 yrs) non-farmers. In contrast, only 38 per cent of better educated (≥6 yrs) non-farmers achieved high coverage.

At the household level, none of the following were related to capsule coverage (after adjusting for distributor

TABLE 3—Adjusted Influence of Village and Distributor Characteristics on Coverage (2 cycles) by Multiple Linear Regression Analysis[a] (n = 189)

Parameter[b,c]	Slope	Std error
Intercept	72.73	5.56
Vill Size 30–59	−3.92	3.68
Vill Size ≥60	−2.65	5.80
Maj Road ≥1 km	12.39	4.21
Cropland none	1.14	3.97
Cropland <½ ha	−14.38	4.78
Occup non-farm	5.52	6.56
Educ > 6 yr	0.40	4.45
Occup * Educ	−18.00[d]	7.90

[a] $r^2 = 0.17$

[b] Reference categories are (in descending order): vill size <30, maj road <1 km, cropland ≥½ ha, occup farmer, educ ≤6 yr.

[c] The "village shop" variable (reported in Table 2) was not included in the regression analyses due to a large number of missing values (n = 42).

[d] Slope estimates effect on coverage of being a non-farmer with >6 yrs education.

performance level): head of household age, sex, education, or occupation; maternal age or education; household size, house construction or its proximity to the center of the village. Similarly, neither sex nor age of children was related to capsule receipt.

Discussion

Questions about community factors which may influence population-based programs are of generic importance for any supplementation, immunization, or other direct delivery health intervention. The Aceh study was a community trial, designed to evaluate the impact of vitamin A distribution on multiple parameters of child health and to identify factors related to coverage within the context of an optimal, single-purpose distribution program which still adhered to existing national guidelines.

Distributor characteristics in this study represented those of individuals freely chosen by village headmen who were given no other formal selection criteria other than a basic ability to read and write the Indonesian language. As selected, slightly more than half of all distributors achieved high coverage at both dosing cycles suggesting a tendency for community leaders, when given the opportunity, to select responsible residents to carry out distribution. This may have been due in part to a clear perception in the community at the outset concerning the program's potential benefits to children. High coverage, however, more closely corresponded to a local lower-middle than upper socioeconomic scale. Less economically developed, more rural villages, and distributors who were either farmers or nonfarmers with minimum education were associated with higher coverage, characteristics more closely aligned with the status quo in this predominantly agricultural economy. Those owning a small patch of cropland performed the poorest, an effect which appeared to be independent of factors measured in the study. Presumably these were distributors engaged in other (unassessed) non-farming means of support. Selection criteria for distributor positions may be more predictive of performance when they represent a balance between the formal skills required for the job (e.g., minimum literacy) and local norms which represent the stable, adult, working class in the population.

Analyses of household and individual child factors revealed no associations with coverage. While this may be

due partly to imprecision of survey-based interviewing, these data suggest that factors other than those measured influence the chances of children receiving vitamin A supplements.

The large dose program in Aceh controlled xerophthalmia,[13] improved growth,[6] and reduced mortality by 34 per cent in target-aged children.[4] Although dramatic, this impact may have been attenuated by the less than complete coverage that prevailed[5]: benefits may have been even greater had all targeted children received vitamin A. Increased attention toward identifying factors that can strengthen community-based programs is likely to lead to both increased effectiveness, efficiency, and sustainability of direct service delivery.

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Learning Disabilities:
Proceedings of 1987 National Conference Available

The Proceedings of the 1987 National Conference on Learning Disabilities is now available from the National Institute of Child Health and Human Development (NICHHD). The 575-page book, available free of charge, provides extensive up-to-date information on what is known about learning disabilities, and also points out what is not known and makes recommendations for future research.

The conference, which was co-sponsored by the Interagency Committee on Learning Disabilities and the Foundation for Children with Learning Disabilities, focused on research findings in five areas: neurobiology of learning and memory; specific developmental disabilities of reading, writing, and mathematics; developmental language disorders; social skills deficits; and hyperactivity/attention deficit disorder. For each topic, the Committee commissioned an expert in the field to write a comprehensive review paper describing recent research related to the cause, diagnosis, treatment and prevention of learning disabilities as well as identifying gaps in knowledge and areas for further research. These documents are published in full in the proceedings, along with commentaries by expert discussants.

Members of the Interagency Committee on Learning Disabilities were drawn from a number of federal agencies, including: NICHHD, the National Institute of Neurological Disorders and Stroke, National Institute of Allergy and Infectious Diseases, National Eye Institute, National Institute of Environmental Health Sciences, and Division of Research Resources of the NIH; the Food and Drug Administration; National Institute of Mental Health; Centers for Disease Control; Environmental Protection Agency; Office of Human Development Services; and Department of Education. NICHHD Director Dr. Duane Alexander was conference chairman. Dr. James F. Kavanagh, Associate Director of NICHHD's Center for Research for Mothers and Children, was executive secretary of the Interagency Committee and co-editor of the proceedings.

To obtain a single free copy of the proceedings, write to NICHHD Publications, P.O. Box 29111, Washington, DC 20040.

Transmission of Enteric Disease Associated with Wastewater Irrigation: A Prospective Epidemiological Study

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Abstract: We conducted a prospective epidemiological study of possible enteric disease transmission by aerosolized pathogens from sprinkler irrigation of partially treated wastewater in 20 kibbutzim (collective agricultural settlements) in Israel between March 1981 and February 1982. Medical data were collected from the patients' files and daily logs of physicians and nurses at each kibbutz clinic (total population 10,231). Episodes of enteric disease were similar in the kibbutzim most exposed to wastewater aerosols (11.6 per 100 person-year) and the kibbutzim not exposed to wastewater in any form (11.0 per 100 person-year). No excess of enteric disease was seen among wastewater contact workers or their families as compared with the unexposed. No negative health effects were detected in this study which involved a large population, including many young children, exposed to treated wastewater aerosols generated at distances of 300–600 m. (*Am J Public Health* 1989; 79:850–852.)

Introduction

Wastewater reuse through irrigation has been a declared a national policy in Israel for the past 30 years. About 250 wastewater irrigation projects currently operate, utilizing some 70 per cent of the wastewater flow of the country; the resulting aerosols expose thousands of farmers and their families who work or live near the irrigated agricultural fields. To see whether dispersed aerosolized pathogens from wastewater sprinkler irrigation systems increase enteric or other diseases, our group has carried out three epidemiological studies in kibbutzim (agricultural collective settlements).

The first study was based on the analysis of official reported cases of communicable disease drawn from the records of the Ministry of Health[1]; the second on an analysis of morbidity as recorded in existing clinic records in the communities under study.[2] The first study, which had a number of serious methodological limitations[3] suggested increases in salmonellosis, shigellosis, typhoid fever, and infectious hepatitis but the second study did not confirm those findings. Nevertheless, the second study found a significant excess of enteric diseases, mainly clinical gastroenteritis, in the 0–4 year old age group in wastewater irrigating kibbutzim during the summer irrigation months only, with no excess of enteric disease detected on an annual basis.

We therefore conducted the present prospective study in 20 agricultural settlements in Israel with a total population of 10,231. The study concerns both occupationally exposed workers and their children and spouses, and the general population residentially exposed to the aerosols formed by sprinkler irrigation with partially treated wastewater. The effluent was not used for the irrigation of directly consumable crops. The presumed routes of exposure to enteric pathogens were dispersion of aerosolized wastewater to nearby residential areas, direct contact with wastewater by irrigation workers and contact of wives and children with wastewater irrigation workers.

Sero-epidemiological studies on the same population groups conducted during the same period are reported elsewhere.[4–9]

Methods

The kibbutz is a voluntary collective community, mainly agricultural, responsible for all needs of members and their families, in which there is no private wealth. Geographic location, climatic factors, quality of water supply, and level of food hygiene vary but socioeconomic and educational levels are uniformly high. The Health Insurance Fund of the General Federation of Labor—Kupat Holim—maintains local health clinics in each kibbutz, according to Western medical standards, with services available to all kibbutz members and residents at no charge. There is a very high level of interpersonal contact between infants and children because the community provides common care from a very young age.

After collecting baseline data in some 100 kibbutzim, 20 were selected on a matched basis for inclusion in the study as representative of Israel kibbutzim in terms of size, location, crops, and population.

All of the wastewater effluent used for irrigation in the 20 kibbutzim was only partially treated—a 5 to 10-day detention period in stabilization ponds—and was generally of poor microbiological quality with coliform concentration of $10^4 - 10^5$/100 ml. The irrigation period of May–August was common to all kibbutzim.

Study Population

The morbidity analysis was based on 20 kibbutzim meeting quality control criteria. The study population (N=10,231) including kibbutz members, candidates for kibbutz membership, their children and parents. The 20 kibbutzim selected were divided into three categories: A) Higher Exposure (seven kibbutzim) wastewater sprinkler irrigation within 300–600 m of residential areas; B) Lower Exposure (six kibbutzim) with wastewater use, but no exposure to aerosols; C) No Exposure (seven kibbutzim) control group; no use of wastewater. The age distribution in the 20 kibbutzim was similar for different categories (see Table 1).

Morbidity Data

Patients' visits to each kibbutz clinic were recorded by nurses or physicians and disease outbreaks recorded by the project staff. The diagnosis or symptoms recorded for each clinic visit were coded according to the International Classification of Health Problems in Primary Care (ICHPPC-2) elaborated by the World Organization of National Associa-

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TABLE 1—Age Distribution in the Study Population of 20 Kibbutzim by Categories (per cent of all ages)

Categories	Age Groups (years)					
	0–5	6–9	10–14	15–24	25–54	55+
Total	16	11	10	9	37	17
A	15	9	11	11	38	16
B	16	9	9	10	38	19
C	16	14	9	7	35	18

tions of General Practitioners and based on the WHO International Classification of Diseases (ICD), 9th Revision. Population estimates for denominator were obtained from the Kibbutz-Movement Statistical Department.

One episode for acute or subacute enteric condition refers to one or more visits to the clinic by the same patient over a certain period of time for the same symptoms or disease. A new episode was recorded when there was an interval of 20 days or more between two visits for the same symptoms and disease. The most common enteric diseases likely to be transmitted by wastewater were chosen for this study (Table 2).

Statistical Analysis

The permutation-randomization test[10] was used to assess the significance of differences in the kibbutz episode rates between categories. This test does not assume that the rates within each kibbutz category follow identical distribution patterns and is not sensitive to the presence of a few outlying kibbutzim within a category.

Results

The morbidity analysis was based on the total enteric conditions reported since the number of specific disease entities were too low to allow for analysis and 95 per cent were non-specific. As Table 3 shows, the overall rates of enteric conditions for all ages in the seven kibbutzim exposed to wastewater aerosols from nearby fields (high exposure) was similar to rates in the six kibbutzim with only occupational exposure to wastewater (lower exposure) and the seven unexposed control kibbutzim (no exposure), both during the year and the irrigation season (May–August).

Table 3 also shows the rates for the 0 to 5-year age group. There were no differences between the higher exposed and no exposure on an annual basis. The apparent excess in mean enteric diseases rates in residentially exposed kibbutzim (49.1/100 vs 36.6/100) during the irrigation period was not statistically significant by the permuation-randomization test (P-value = 0.22) and was due to the high rate in one exposed kibbutz. When that kibbutz was excluded the mean rate for

TABLE 2—Number of Episodes of Selected Enteric Conditions Reported during Study

Salmonellosis	6
Amebiasis	1
Campylobacterosis	5
Diarrhea	512
Shigellosis	17
Giadiasis	9
Viral hepatitis	13
Gastroenteritis, clinical	539
Total Enteric Episodes	1102

TABLE 3—Episode Rate of Enteric Conditions (per 100 person-years) by Kibbutz Categories (all ages and 0–5, both sexes)

Age Group	Category	(No.)	Population	Episodes			
				Year*		Irrigation Period	
				N	Rate	N	Rate**
All Ages	Higher Exposure	(7)	3,562	412	11.6	220	18.5
	Lower Exposure	(6)	2,684	251	9.4	135	15.1
	No Exposure	(7)	3,985	439	11.0	229	17.2
	Total		10,231	1102	10.8	602	17.7
0–5	Higher Exposure	(7)	538	140	26.0	88	49.1
	Lower Exposure	(6)	436	87	20.0	52	35.1
	No Exposure	(7)	656	173	26.4	80	36.6
	Total		1,630	400	24.5	220	40.5

*From March 1981 to February 1982
**Four months (May–August 1981) episodes N×3×100/p = rate/year

the high exposure group was reduced to 29.4/100, which is lower than the no exposure group.

No excess of enteric disease outbreaks was found in wastewater irrigating kibbutzim, neither was there any association between drinking water quality and morbidity. No excess of enteric disease was found among wastewater contact workers, their wives or children.

Potential confounding factors, including microbial quality of drinking water and irrigation water, occupation, geographical region, age, and sex were not found to be consistently related to enteric disease rates.

Discussion

In our studies, low levels of bacteria and viruses of wastewater origin have been detected at distances of 730 m downwind of wastewater irrigated fields.[11] Although aerosolized enteric pathogens certainly can cause infection and disease in susceptible individuals, these morbidity data indicate that this does not commonly occur in the exposed study population.

These results contradict those of our first retrospective study,[1] which suggested possible transmission of typhoid, salmonellosis, shigellosis, and infectious hepatitis. It is possible that inaccuracies in official published figures which were used for the analysis account for the different results.[3]

Wastewater irrigation workers, the most highly exposed group, may be protected by their high levels of immunity to a broad spectrum of common enteroviruses.[9] Similar findings have been reported among wastewater workers in the United States, where initial levels of herd immunity are lower.[12]

Susceptible infants and children would be the most likely to show ill effects, as our earlier retrospective study[2] suggested, but no clear effects were found in the current study. During the irrigation season, young children in one kibbutz exposed to wastewater aerosols had high enteric disease rates, but children in the other six exposed kibbutzim did not.

The morbidity data from the present study showed no excess of enteric disease in all ages or in the 0 to 5-year age group on a yearly basis in wastewater aerosol exposed communities as compared to the non-exposed control communities. This finding is identical to that of the second

retrospective study.[2] This congruence of results between two independent studies reinforces this key finding.

A possible limitation in the study is that the high level of person to person contact in the communal infant and children homes and the common dining rooms could mask infrequent occurrence of disease associated with wastewater sprinkler irrigation. A strength of the study is that it included many infants with low immunity exposed to wastewater aerosols of poor microbial quality at relatively close range. If enteric diseases can be transmitted by such exposure to a significant degree, it should have been detectable under the conditions of this study. However, the evidence provided by this study does not point to any significant health risk associated with exposure to treated wastewater aerosols generated at distances of some 300–600 m from residential areas.

ACKNOWLEDGMENTS

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NIMH Earmarks $4.5 Million in 1989 for Homeless Mentally Ill

A new federal effort to benefit the homeless mentally ill was announced recently by Lewis L. Judd, MD, director of the National Institute of Mental Health (NIMH). "Improving the care and treatment of mentally ill individuals who are homeless is among the most pressing public health needs facing this country, and one of the most important priorities of the Institute," Dr. Judd said.

NIMH has budgeted some $4.5 million in Fiscal Year 1989 to launch a new Institute-wide program of research grants and demonstration projects that focus on the homeless mentally ill. The program will obtain essential knowledge about this population, their needs, and the best strategies for their care and treatment. Approaches to rehabilitative and other services for the homeless mentally ill will be rigorously tested and evaluated in actual service settings. The Institute also created a new unit in the NIMH director's office that will spearhead efforts to improve services to the homeless mentally ill.

Studies have shown that 30–35 percent of homeless adults suffer from a severe mental illness such as schizophrenia or manic-depressive illness. Less is known about the extent of mental illness among homeless families with children, a growing subgroup among the homeless. NIMH has supported research to identify the mental health, housing, and other support needs of the homeless mentally ill, and has also funded research, demonstration projects, and technical assistance efforts focusing on homelessness and its relation to mental illness. The new initiative represents a major expansion of those efforts.

The newly created NIMH Office of Programs for the Homeless Mentally Ill, to be located at NIMH in the Parklawn Building, 5600 Fishers Lane, Rockville, MD 20857, will be directed by Irene S. Levine, PhD; Debra J. Rog, PhD, will be associate director.

Accidental Ingestions of Oral Prescription Drugs: A Multicenter Survey

BARBARA J. JACOBSON, AMY R. ROCK, MPH, MURRAY S. COHN, PHD, AND TOBY LITOVITZ, MD

Abstract: Accidental ingestion of oral prescription drugs by children under age five continues to be a public health problem. A telephone survey of 1,866 ingestion incidents reported to nine poison control centers was conducted in the spring of 1986. Accidental ingestion occurred most often with a two-year-old child (42 per cent) in their own home (82 per cent). Thirty-five per cent of the toxic drugs ingested at home belonged to someone other than the immediate family, most often a grandparent. Toxic drugs were more often out of their usual storage location and in non-child-resistant prescription packaging, a nonprescription container, or in no container. Twenty-two per cent of all child-resistant packages were non-functional. Overall, at least 61 per cent of all medications had no child-resistant barrier at the time of ingestion. Accessible storage locations such as the kitchen counter, table top, or top of a dresser or bedside table greatly increased the risk of accidental ingestion. These results suggest the need for effective child-resistant packaging that is easier for all adults to use. (*Am J Public Health* 1989; 79:853–856.)

Introduction

The US Consumer Product Safety Commission (CPSC) is responsible for protecting children under the age of five from serious personal injury and death under the authority of the Poison Prevention Packaging Act of 1970. Its 1974 regulations require child-resistant packaging for oral prescription drugs. Previous studies have established that child-resistant packaging has indeed played a role in reducing injuries and deaths due to accidental ingestions.[1–3] Nevertheless, childhood poisoning is often due to ingested medications.[4–6]

In 1986, there were 49,281 accidental ingestions of oral prescription drugs involving children under age five reported to the American Association of Poison Control Centers (AAPCC), National Data Collection System;[7] 131 involved life-threatening symptoms, disfigurement, or residual disability; and four resulted in death. These numbers underrepresent the actual number of ingestions because they do not include ingestions reported only to other health care professionals and facilities.

Even without serious injury or death, ingestion incidents cause obvious emotional trauma to the child and his/her family. Therefore, the CPSC conducted a retrospective study to characterize specific factors that contribute to accidental oral prescription drug ingestions involving children under five years of age.

Methods

The study was designed to provide descriptive data about 2 025 incidents (225 from each of nine poison centers) of oral prescription drug ingestions involving children under 5 years of age.

One poison center from each of the nine US Census regions was selected to administer the study questionnaire and report to the American Association of Poison Control Centers (AAPCC) National Data Collection System.*

Interviewers from each poison center, selected by the center director, were (with one exception) poison center specialists who normally handle incoming calls at the center; they were trained in case selection and uniform interviewing techniques.

The study began in February 1986 and continued until the target of at least 225 cases of suspected prescription drug ingestions had been recorded by each interviewer. Study cases were consecutively selected from all incoming calls on the basis that the accidental ingestion involved an oral prescription drug and a child under five years of age. Excluded were ingestion incidents involving veterinary drugs; prescription ointments and creams in tubes; eye drops and ear drops; and lotions for dermal application. Also excluded were incidents where an adult gave a child multiple doses or too large a dose of prescription medication.

Data about the child, the medication ingested, and the medical effects experienced by the child were routinely recorded on the report form.

The study questionnaire was completed during a follow-up telephone interview with the individual who initially called the poison center or another adult who was in the household or at the location at the time of the ingestion incident. If necessary, at least 10 attempts to reach a respondent were made, including two in the evening and two on the weekend. The interviewer made the first attempt to contact a respondent within two working days of the initial incoming call. The involvement of an oral prescription drug and a child under age five was verified at the beginning of the interview.

The questionnaire was designed to collect information about the child, the ingestion site, the intended user of the medication, identification of any packaging, and the specific location of the medication when the child found it. For incidents involving non-child-resistant packaging, respondents were asked whether a specific request was made at the pharmacy for a non-child-resistant package.

All respondents who still had the involved package were asked if they would be willing to mail it to us at CPSC. Respondents who agreed were sent a postage-paid, preaddressed mailing bag.

A questionnaire was recorded for each child and for each drug involved in an incident. For incidents involving more than one child and/or more than one drug, the data were weighted to a one child/one drug basis. Any resulting fractional results were rounded to the nearest whole number. The results are descriptive and are presented on a per child or per drug basis as appropriate.

Address reprint requests to Barbara J. Jacobson, US Consumer Product Safety Commission, Directorate for Health Sciences, Room 724, Washington, DC 20207. Ms. Rock and Dr. Cohn are also with US CPSC; Dr. Litovitz is with the National Capital Poison Center at Georgetown University Hospital, Washington, DC. This paper, submitted to the Journal May 18, 1987, was revised and accepted for publication October 4, 1988. This study, conducted at the expense of the US Government, is considered in the public domain.

*San Diego Regional Poison Center, San Diego, CA; National Capital Poison Center, Washington, DC; Kentucky Regional Poison Center, Louisville, KY; Louisiana Regional Poison Center, Shreveport, LA; Massachusetts Poison Control System, Boston, MA; Children's Hospital of Michigan Poison Control Center, Detroit, MI; Hennepin Poison Center, Minneapolis, MN; Pittsburgh Poison Center, Pittsburgh, PA; Intermountain Regional Poison Control Center, Salt Lake City, UT.

Results

There were 1,953 ingestion incidents involving 2,015 children under age five and 1,982 drugs: 1,866 incidents involved one child and one drug, 60 incidents involved two children and one drug, 24 incidents involved one child and two drugs, two incidents involved one child and three drugs, and one incident involved two children and two drugs. All incidents involved at least one oral prescription drug.

The study incidents were approximately 4 per cent of the reported 49,281 oral prescription ingestions involving children under age five in 1986.

The 1,953 incidents are 92 per cent of the 2,128 eligible incidents reported to the centers; 175 incidents were not included: 63 incomplete questionnaires, 44 refusals, 29 no eligible respondent, 25 wrong number given, and 14 non-working numbers given.

Eighty-nine per cent of the respondents were the individual who initially reported the ingestion to the poison center, 87 per cent were the child's parent, most often his/her mother.

The child's household most frequently had four members with a total annual income of $20,000 to $50,000. Eighty-one per cent of the children involved were White.

The drug categories included:

- *Antimicrobials*, including 297 flavored liquid antibiotics prescribed for the child or his/her sibling, were the most frequently ingested drugs (Table 1).
- *Oral contraceptives*, most often prescribed for the child's mother.
- *Toxic drugs*, most frequently prescribed for the child's grandmother (38 per cent, n = 228) or mother (32 per cent, n = 192).

Overall, more males were involved (56 per cent) than females (44 per cent). The peak age was two years (42 per cent); few incidents involved children under age one. Two-year-olds ingested non-toxic drugs as frequently as toxic drugs (41 per cent vs 44 per cent). Children under age two ingested more toxic than non-toxic drugs (36 per cent vs 19 per cent); children over age two ingested more non-toxic than toxic drugs (37 per cent vs 23 per cent).

TABLE 1—Categories of Oral Prescription Drugs Ingested by 2015 Children under Age Five

Drug Categories	n = 1982 (%)
Non-toxic*	
Antimicrobials	516 (26)
Oral contraceptives	200 (10)
Hormones	87 (4)
Toxic**	
Analgesics	199 (10)
Cardiovascular drugs	174 (9)
Sedatives/hypnotics, antipsychotics	124 (6)
Asthma therapies	41 (2)
Anticonvulsants	33 (2)
Antidepressants	30 (2)
Other	
Electrolytes/minerals	124 (6)
Cough and cold preparations	104 (5)
Vitamins	90 (5)
Diuretics	50 (3)
Anticholinergics	38 (2)
Antihistamines	35 (2)
Gastrointestinal preparations	30 (2)
Muscle relaxants	19 (<1)
Miscellaneous	88 (4)

*Drugs rarely associated wtih adverse effects other than minor gastrointestinal symptoms, even when ingested in large amounts.

**Drugs commonly associated with severe systemic toxicity.

Site and Intended User

Eighty-two per cent (n = 1652) of all children were at home when the ingestion occurred. The intended user of the medication was most often the child's mother (27 per cent) followed by the child him/herself (22 per cent). Children that were not at home were most frequently at the home of a grandparent (60 per cent) or an aunt or uncle (14 per cent). Thirty-five per cent of the toxic drugs ingested in the child's own home belonged to someone not in the child's immediate family (Table 2), most frequently the child's grandparent.

More toxic than non-toxic drugs were ingested at locations outside the child's home (32 per cent vs 9 per cent).

Of the 2,015 children, 1,471 developed no symptoms, 247 were estimated to have a non-toxic exposure, and four exhibited unrelated symptoms; 231 children experienced minor symptoms, 18 experienced moderate symptoms usually requiring treatment, 37 were estimated to have a potentially toxic exposure, and seven experienced life-threatening symptoms. There were no deaths.

Over half of all drugs ingested were located in the kitchen/dining area (53 per cent). Liquid flavored antibiotics were the most frequently ingested preparation and the refrigerator was the most common location (13 per cent). Twelve per cent of the ingested drugs were on the kitchen counter and 7 per cent were on the table.

The toxic drugs were found in the bedroom almost as often as in the kitchen (31 vs 38 per cent). In the bedroom, medicine was most frequently stored on top of the dresser or bedside table.

Only 2 per cent of all drugs were stored in a bathroom medicine cabinet. Ten per cent were stored in a purse (30 per cent of all birth control pills).

Overall, the drugs were stored in their usual location only 53 per cent of the time. In fact, the toxic drugs were more often out of their usual location (65 per cent of the time) and were frequently on the floor or a low, open storage place.

Table 3 shows that more toxic than non-toxic drugs were in non-child-resistant prescription packaging (18 vs 7 per cent).

Seventy-five per cent (n = 235) of all non-child-resistant packages were reportedly dispensed by the pharmacist without a request from the purchaser or the prescribing physician.

Most of non-child-resistant packages (76 per cent) had been dispensed to parents of a child under age five.

TABLE 2—Site Where Accidental Ingestion Occurred and Intended User of the Most Potentially Toxic and Non-toxic Drugs

Site and Intended User	Toxic Drugs n = 601 %	Non-toxic Drugs n = 803 %
Child's Home	68	91
Immediate Family*	(41)	(82)
Other Live-ins	(14)	(4)
Visitor	(10)	(4)
Intended User Unknown	(3)	(1)
Some Place Else	32	9
Immediate Family*	(2)	(3)
Not Immediate Family	(30)	(6)
Intended User Unknown	(<1)	(0)

*Includes child, siblings, and parents.

TABLE 3—Type of Package Drug Was Stored In at the Time of Ingestion

Variables	Toxic n = 601 %	Non-toxic n = 803 %	All Drugs n = 1982 %
Child-resistant*	29	47	39
Non-child-resistant	70	27	50
Child-resistant package (closure left off or loose)	(6)	(9)	(9)
Non–child-resistant package	(18)	(7)	(16)
No container (loose pills)	(29)	(9)	(17)
Transferred to another container	(17)	(2)	(8)
Unit dose (oral contraceptives)**	0	25	10
Unknown	1	1	<1

*Includes child-resistant unit dose packaging.
**Excluded from child-resistant packaging requirement.

We received 382 of the requested prescription containers, 31 per cent of the 1,228 reportedly still available. Based on the opening and closing forces measured on mechanical testing equipment, 200 of the 306 (65 per cent) child-resistant packages received were not functioning within the normal force range for their specific design type. If the 65 per cent non-functioning packages received were extrapolated to all child-resistant containers (excluding unit-dose), the percentage of medications with no child-resistant barrier could be higher than 70 per cent.

The majority of the continuous threaded closures that were not functioning had been affected by the sticky liquid antibiotic in the package. Closures on the lug containers were non-functional due to wear associated with use.

Diseussion

This study had several limitations that may have biased the results:

- These data refer only to telephone calls to poison centers. We made no effort to gather data about ingestions reported only to other sources. Such ingestions could represent some of the most potentially toxic exposures that required immediate treatment in hospital emergency rooms.
- Our attempt to validate the package identification and measure function was incomplete because only 31 per cent of the prescription packages were sent to us. Some of the packages were no longer available at the time of the interview, which contributed to the low response rate. There is no reason, however, to suspect that the package identification is incorrect because those packages received did confirm the identification. Results on function are incomplete but do indicate the need for further evaluation of prescription packages.
- The reliability of responses to sensitive questions was not documented. For example, child-resistant closures that were left off or loose, drugs transferred out of their original prescription container, and drugs out of their usual storage place may be underreported because the truth involves an admission that carelessness contributed to the accidental ingestion.

A pharmacist who dispenses non-child-resistant packaging without a specific request from the consumer or his/her physician violates the provisions of the Poison Prevention Packaging Act. Respondents who were not the intended user of the ingested drug reported that a child-resistant package had been requested more frequently than respondents who were the intended user of the drug. This discrepancy may be related to the sensitivity of the question. Therefore, the results cannot be used to reflect pharmacist dispensing practices.

Other studies have shown that the personality and age-specific behavior characteristics of young children are important risk factors in accidental poisoning.[8,9] Children two to three years old were shown to be at greatest risk, confirming the results of several epidemiological studies.[4,6] The results of this study consistently showed that a lack of an effective child-resistant packaging barrier contributes to accidental ingestions of oral prescription medications. The accessible kitchen storage location identified is well documented in other studies and increases the risk of ingestion.[10–12] Storage on the kitchen counter, table top, or top of a dresser or bedside table suggests that the medication was being stored where it was convenient for the user to take. Results of a previous study found that the majority of ingested products were in use within 24 hours before the accident.[13]

A grandparent's home as an important risk factor has been identified in other studies as well.[14,15] Our results indicate that a grandparent's medication presents increased risks due to toxicity and the likelihood that the medicine will have no child-resistant barrier. A previous prospective study of poisonings that occurred away from home concluded that the products involved, most frequently drugs, were more potentially toxic and that these environments were seldom adequately prepared to handle toxic exposures.[16]

As the elderly population continues to expand and age, increased attention must be directed to educate these individuals about the risk of ingestion from their oral prescription drugs.

It is clear that packaging manufacturers have a responsibility to produce effective child-resistant prescription containers that are easier for adults of all ages to use.

Additionally, some practical method of safe storage is needed for drugs, particularly while they are in use.

Pharmacists, as well as dispensing physicians, have a key educational role to play when dispensing prescriptions, especially toxic drugs dispensed in non-child-resistant packaging.

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A summary of the results of this study was presented at the annual scientific meeting of the American Association of Poison Control Centers, American Academy of Clinical Toxicology, American Board of Medical Toxicology, and Canadian Association of Poison Control Centers, Sante Fe, New Mexico, September 25–30, 1986.

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Congressional Task Force on Tobacco and Health

A new Congressional Task Force on Tobacco and Health—comprised initially of a bipartisan group of 35 members of the US House of Representatives—was announced in early May to focus congressional attention on the growing public policy issues associated with tobacco use.

Co-chairs of the task force are Rep. Richard J. Durbin (D-IL) who originated the 1987 law banning smoking on commercial airline flights of two hours or less duration, and Rep. Bob Whittaker (R-KS), sponsor of legislation to prohibit the sale of tobacco to minors and further regulate the sale and promotion of tobacco products. Among the other members are Reps. Henry Waxman (D-CA), chair of the House Energy and Commerce Subcommittee on Health, and Fortney Pete Stark (D-CA), chair of the House Ways and Means Subcommittee on Health.

Reps. Durbin and Whittaker said that the task force evolved as a result of the Conference on Tobacco Use in America, held last January at the M. D. Anderson Cancer Center in Houston, Texas. That conference brought together physicians, public health officials, and members of Congress specifically to address tobacco-related health issues. A comprehensive strategy for reducing tobacco use in America, drafted as a result of that conference, is to be announced by the task force. The stated purposes of the Task Force on Tobacco and Health are to:

- Educate Members of Congress and the public on tobacco use and its health effects;
- Build coalitions with other congressional organizations and public groups interested in tobacco use and tobacco-related health issues;
- Develop and coordinate strategies for advancing tobacco-related health legislation; and
- Assist Members of Congress in addressing the problems associated with tobacco use.

Individual members of the task force have sponsored a wide variety of bills aimed at helping to implement the goal of a smoke-free society by the year 2000. Those bills include: expansion of smoke-free zones in public buildings and transit systems; limitations on tobacco advertising; strengthening the warnings on cigarette labels; and raising the federal excise tax on cigarettes.

For further information, contact Steve Blakely in Rep. Durbin's office, 129 CHOB, 202/225-5271, or Kelly Presta in Rep. Whittaker's office, 332 CHOB, 202/225-3911, Washington, DC 20515.

The Power of Policy: The Relationship of Smoking Policy to Adolescent Smoking

Mary Ann Pentz, PhD, Bonnie R. Brannon, EdD, Ventura L. Charlin, PhD, Edith J. Barrett, PhD, David P. MacKinnon, PhD, and Brian R. Flay, DPhil

Abstract: This study examined the effects of smoking policy on 4,807 adolescents in 23 schools over a two-county area in California. Amounts and prevalence rates of adolescent smoking were measured with a self-report survey and a biochemical measure; school smoking policy was measured with two independent surveys of school staff. Policy effects were evaluated with multiple and logistic regression analyses controlling for school-level socioeconomic status and environmental support for teaching and administration. Of the 23 schools, 100 percent had a formal written and regularly enforced policy component restricting student smoking on school grounds, 94 percent restricted students leaving school grounds, 65 percent restricted smoking near school grounds, and 57 percent had a smoking prevention education plan. Schools with policies having all four versus less than four components, high versus low emphasis on prevention, and a low versus high emphasis on cessation reported lower amounts of smoking in the last week and in the last 24 hours. Punitive consequences of policy violation had no effect. Results were compared to school staff observations of adolescent smoking, and school archival records of student smoking violations in the last year. Results suggest that school smoking policy is associated with decreased amounts of smoking in adolescents. (*Am J Public Health* 1989; 79:857–862.)

Introduction

Despite widespread knowledge of the health risks of cigarette smoking, 26.5 percent of adults and 18.7 percent of graduating high school seniors in the United States continue to smoke daily.[1,2] Fortunately, among adults, the prevalence of smoking and the associated rates of chronic disease mortality rates have declined steadily in the last 20 years.[1,3] However, among adolescents, smoking has decreased only 1.6 percent in the last five years.[4]

The effectiveness of school-based educational programs for preventing early adolescent smoking has been well documented.[5,6] Nevertheless, additional prevention strategies may be required to produce long-term changes in adolescent smoking behavior.

Over 90 percent of adult smokers who quit smoking do so on their own.[3,7–9] This percentage suggests that the steady decline in adult smoking prevalence rates may be at least partially attributable to efforts that have been aimed at motivating smokers to quit, such as anti-smoking mass media campaigns, increased taxation of cigarettes, public smoking education, and regulatory policies that restrict smoking in the workplace or public settings.[10–15]

No studies have yet evaluated the effectiveness of policy in reducing adolescent smoking. A logical approach would be to evaluate the effects of school smoking policy on adolescents, controlling for school-level socioeconomic status (SES), support for teaching and administration, and variables differentiating components of policy that relate to educational programming for smoking prevention from components that relate to regulation of smoking.[16,17] We undertook such a study.

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Methods

Subjects

Subjects included 4,807 seventh grade students, 104 teachers, 21 principals, and 23 school clerks from 23 middle/junior high schools in Los Angeles and San Diego counties. The 23 schools represented all of the middle/junior high schools from three large school districts that are participating in a five-year trial of the effects of varied smoking education programs on prevention of adolescent smoking.[18] The average grade size per school was 213, with a student population characterized as high minority and low socioeconomic status (20 percent White, 27 percent Black, 44 percent Hispanic; 23 percent receiving free lunch as part of the federal Aid to Families with Dependent Children (AFDC) program). The population was 49 percent female.

Informed consent was required of both student and parent. Of the 6,101 students enrolled in seventh grade at the time of the study, 319 declined to participate and 975 were absent, resulting in a response rate of 79 percent. Of the school staff solicited for participation, 100 percent of teachers, 91.3 percent of principals, and 100 percent of school clerks responded and participated in assessment. Two principals (8.7 percent) who initially agreed to participate were unavailable during the data collection period and, consequently, were not assessed.

Policy

The California State Education Code bans all student smoking on school grounds in middle and junior high schools. At the initiation of this study in 1986, the policy had been in effect for an average of 34 years in the participating middle and junior high schools, and students had been exposed to the policy for an average of 13 months.

Measures

Sources of data included school archival records, a student self-report survey and biochemical measure of smoking, and surveys of teachers, principals, and school clerks.[19] (See Appendix for details of specific items analyzed for this study.)

Students completed a 109-item questionnaire measuring amount (number of cigarettes smoked) and prevalence rates (percent of smokers) of cigarette smoking in lifetime, last

week, and last 24 hours; related attitudes and behaviors; and demographic characteristics (gender, race/ethnicity, socioeconomic status).[20] During questionnaire administration, students also provided an expired air sample for carbon monoxide analysis of smoking, and were informed of the use of the sample as a test for smoking.[21,22]

The science and health education teachers and the principal of each school independently completed a 96-item questionnaire about staff and student smoking policies, including restriction of smoking on and near school grounds, a closed campus limiting opportunities for smoking off school grounds, and a formal educational plan for smoking prevention.[16,23] Staff awareness, perceived effectiveness of policy, and observations of staff and student smoking were also assessed. Environmental support for teaching and administration was measured using a modified form of the Moos Classroom Environment Scale, which was originally designed to assess supportive teaching environment in the classroom as it relates to student academic and social behavior.[24]

School clerks were interviewed with a 21-item survey which measured the extent to which smoking was allowed on school grounds, the length of years policy had been in effect, level of enforcement of policy, change in the number of smoking violations in the school in the last year, consequences of policy violation, and emphasis of the policy on punishment, prevention, and cessation.

Statistical Analysis

Separate t-tests were conducted to determine the equivalence of principal and teacher responses for pooling these groups in analyses. Principals and teachers did not differ in their reports of smoking levels of students or staff, or existence of policy or smoking prevention programs (none of the schools had smoking prevention programs). Within-school internal consistency of responses was high (calculated as a Cronbach alpha of all items across respondents within each school, average alpha = .83, ranging from .50–.99 for each school). Thus, principal and teacher ratings were averaged to produce one rating for each item for each school. For four schools in which the principal survey was incomplete, the rating for each item for that school was based on the average rating by the four to five teacher respondents.

Multiple linear and logistic regression analyses were conducted on each adolescent smoking item (amount and prevalence), with each policy variable (comprehensiveness, prevention emphasis, cessation emphasis, punishment emphasis) entered as an independent variable, and school-level socioeconomic status and school environmental support entered as covariates. To increase test validity, staff awareness and perceived effectiveness of policy were additionally controlled for in analyses of the effect of number of policy components. Separate t-tests were also conducted to determine whether individual policy components had a significant effect on smoking. The SAS software program was used for all analyses (GLM ANOVA for linear regression; CATMOD for logistic regression.[25] Analyses were conducted using school as the unit by aggregating individual respondents' data by school, and then merging the three data sets by a three-digit school identification code.

Results

Schools varied widely in the number and combination of policy components related to regulation of smoking (no smoking on school grounds, near school grounds, and closed campus), but were about evenly divided on the component related to smoking education (smoking prevention education plan). Characteristics of schools with and without a smoking prevention education plan were similar (Table 1). The groups did not differ significantly on any of the characteristics.

Averaged across all schools, 5.1 percent of adolescents reported having smoked in the last week, and 2 percent in the last 24 hours. Smokers averaged 7.1 cigarettes in the last week, and 2.3 cigarettes in the last 24 hours.

In a separate sub-study conducted under more controlled field test conditions, number of cigarettes smoked in the last 24 hours was assessed on a random sample of the same students one year later (n = 2,500) and correlated with a double expired air measurement following procedures by Biglan, *et al.*[21] The correlation of self-reported smoking with parts per million concentrations of carbon monoxide averaged r = .41, consistent with the average correlation of .35 reported by other researchers for similar age/grade groups.[22]

Effect of Policy on Smoking

As expected, all of the 23 schools had a formal rule banning smoking on school grounds; 22 (94 percent) also had a closed campus. However, only 65 percent had a rule banning smoking near school grounds, and 57 percent had a formal health education plan for smoking prevention programming, although none of the schools had implemented smoking prevention education at the time of this study. All of the school clerks reported that smoking policy was enforced "very regularly."

Table 2 shows the relation of policy to adolescent smoking, dichotomizing schools into two groups formed by the median split of each variable, for ease of illustration. Schools with all four policy components had lower smoking prevalence rates and lower mean amounts of smoking per smoker in the last week and in previous 24 hours than schools with less than four components. When schools with all four components were compared to schools with three or two components, differences in amount of smoking were more pronounced (see Figure 1). Regression analyses—using all four policy components and controlling for socioeconomic status, school environmental support, and staff awareness

TABLE 1—Characteristics of Schools with and without a Smoking Prevention Education Plan in Smoking Policy

Characteristics	Without Smoking Prevention Education Plan (n = 10)	With Smoking Prevention Education Plan (n = 13)
Mean Grade Size	211	214
% Female	49	49
% Minority	83	78
% English 2nd Language	19	10
% Aid to Families with Dependent Children (AFDC)	26	21
Mean California Achievement Profile (CAP)	142	149
% Fathers in Professional Occupations (School SES)	20	23
Mean Community SES	3.09	2.86
Mean School Support†	.01	.04
% Adolescents Smoked in Lifetime	41	41
% Staff Currently Smoking	9	12

NOTE: Based on N = 23 schools, N = 4,807 adolescents, N = 125 school staff; school as unit of analysis. All comparisons were non-significant, based on t-tests for mean values and tests between proportions for percentage values, % and means unadjusted.
†Standardized score (x = 0, SD = 1, range = −.95 to .89).

TABLE 2—Prevalence and Amount of Adolescent Smoking by Level of Smoking Policy

Policy Variables	Prevalence of Smoking (% of Smokers)		Amount of Smoking (Mean No. of Cigarettes per Smoker)	
	Last Week	Last 24 Hours	Last Week	Last 24 Hours
Number of Policy Components				
Four	4.93	1.84	3.87	1.81
Less than four	5.60	2.43	6.05	1.99
Difference	−.67	−.59	−2.18	−.18
(90% CI of difference)†	(−2.52,1.18)	(−1.46,.28)	(−5.29,1.26)	(−1.01,1.37)
Prevention Emphasis				
High	4.31	1.88	3.11	1.46
Low	5.77	2.15	6.46	3.43
Difference	−1.46	−.27	−3.35	−1.97
(90% CI of difference)	(−3.01,.09)	(−1.06,.52)	(.28,6.40)	(.69,3.11)
Cessation Emphasis				
High	5.29	2.18	5.87	2.55
Low	4.72	1.78	2.92	1.63
Difference	−.57	.40	2.95	.92
(90% CI of difference)	(−2.30,1.14)	(−1.23,.43)	(−3.91,−1.83)	(−1.75,−.09)
Punishment Emphasis				
High	4.91	1.92	4.57	2.16
Low	5.38	2.21	5.00	2.53
Difference	−.47	−.29	−.43	−.37
(90% CI of difference)	(−1.28,2.23)	(−.56,1.13)	(−3.12,3.97)	(−1.03,.33)

Based on N = 23 schools, school as unit of analysis. Means are adjusted for SES and school environmental support for teaching and administration, for the policy component variable, means are also adjusted for staff awareness and perceived effectiveness of policy.
†CI = Confidence Intervals

and perceived effectiveness of policy—indicated that number of policy components was related to lower amounts of adolescent smoking. Prevalence rates of smoking and smoking violations reported by school clerks were less strongly related to number of policy components.

Effect of Policy Emphasis

When policy emphasis variables were regressed on each smoking variable, controlling for socioeconomic status and school environmental support, results indicated that a high emphasis on prevention and a low emphasis on cessation were associated with lower amounts of smoking in the last week and 24 hours (see Table 2 and Figure 2); results were similar for smoking in the last week and the last 24 hours. Schools with a high emphasis on prevention also had lower weekly smoking prevalence rates and decreased school smoking violations in the last year. When prevention and cessation emphasis variables were entered simultaneously into regression analyses, both had an impact on amount and prevalence rates of smoking. However, punishment and severity of consequences for violation of smoking (number of consequences) had no effect. In all analyses, higher SES was marginally related to lower amounts of smoking.[*]

In general, the emphasis of policy on assisting (prevention and cessation) versus punishing students was significantly related to lower amounts of adolescent smoking and, less consistently, to lower smoking prevalence rates.

Discussion

Results of the study indicate that the number of components in school smoking policy (presumably representative of policy comprehensiveness) has a marginal effect on adolescent smoking, and the use of punitive measures to regulate smoking has no effect. On the other hand, the relative emphasis of policy on educating adolescents to stop smoking, and particularly to prevent smoking, rather than regulating smoking by punishment, is significantly related to lower recent smoking. More consistent effects were obtained on amounts of smoking rather than on prevalence rates, suggesting that current smoking policies for youth may work somewhat better for decreasing levels of smoking rather than for deterring smoking altogether or promoting cessation. Results do not appear to be confounded by any historical differences among schools in smoking risk or need for smoking cessation, as evidenced by the lack of difference among schools in school demographic characteristics, lifetime smoking prevalence rates of students, and staff smoking rates.

Our results are consistent with other recent cross-sectional studies of the relation of public or worksite no-smoking policy to smoking in adults, which suggest that policy may decrease cigarette consumption in settings affected by the policy, but not prevalence rates.[26–29] However, none of these studies is directly comparable to the present study, since each was based on a single policy with no differentiation of components or control for socioeconomic and environmental support variables; subsampling was used, with lower response rates; and policy effect was evaluated on an indirect or a single measure of smoking.

Adolescents may be more likely to comply with a smoking policy that is supportive and positive in focus. Such an interpretation is consistent with the findings of recent school-based smoking prevention studies that have shown positive skill-building prevention programs to be more effective in preventing or delaying the onset of smoking than programs focused on negative physiological consequences of smoking.[6] Also noteworthy is the finding that although all schools complied with the California state law banning smoking on school grounds, schools varied widely in representation of other policy components, staff awareness of policy, and consequences of policy violation. Perhaps most

[*]Data available on request to first author.

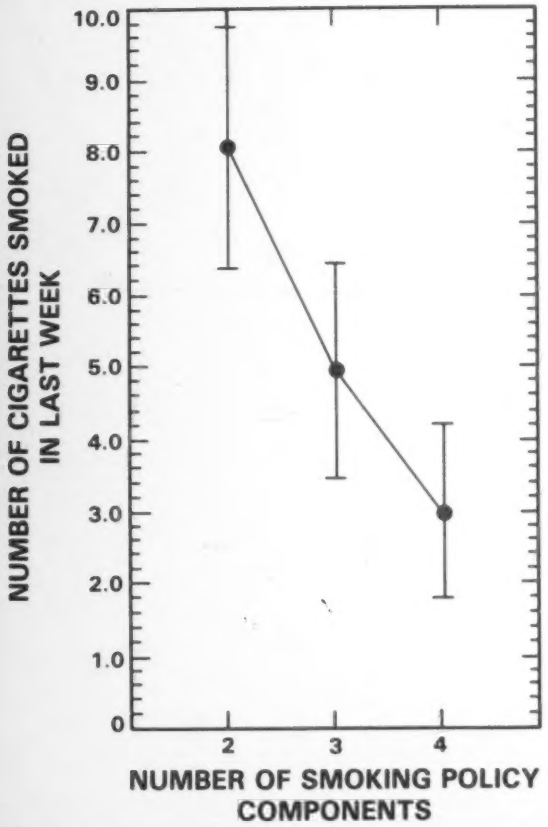

FIGURE 1—Amount of Smoking (no. of cigarettes smoked per smoker) in the Last Week in Schools with Two (n = 6), Three (n = 7), or Four (n = 10) Smoking Policy Components
Means are adjusted for school level SES, support for teaching and administration, and staff awareness and perceived effectiveness of policy.

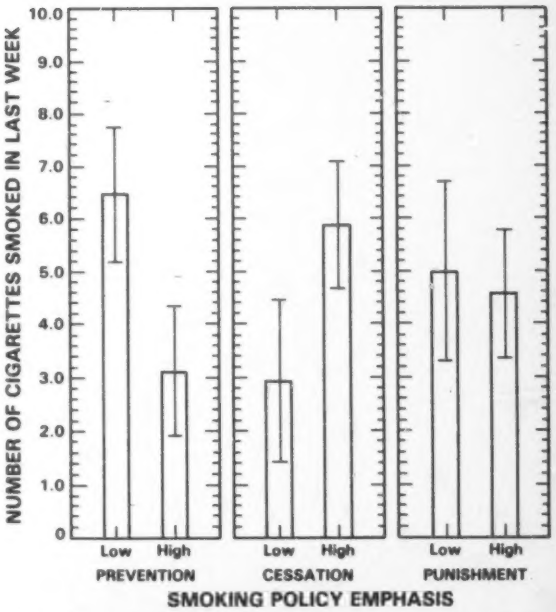

FIGURE 2—Amount of Smoking (no. of cigarettes smoked per smoker) in the Last Week in Schools with High and Low Emphasis on Smoking Policy
NOTE: Prevention Emphasis: n = 11 low, 12 high
Cessation Emphasis: n = 9 low, 14 high
Punishment Emphasis: n = 8 low, 15 high
Means are adjusted for school level SES, and support for teaching and administration.

surprising was the finding that none of the schools had implemented a smoking prevention program, despite reports that 57 percent had a formal educational plan for smoking prevention. Simply having a formal statement of policy is not likely to impact on adolescent smoking—the policy should emphasize smoking prevention and be actively implemented.

The findings of this study must be evaluated in the context of design limitations of policy studies. Although a longitudinal relationship between policy and smoking is assumed, since the smoking policies had been in effect in schools an average of 34 years and adolescents had been exposed to these policies in middle school an average of 13 months before smoking was measured, the study is cross-sectional. An experimental design of longitudinal policy effect would require random assignment of schools or other units to policy or no-policy conditions, after which policy would be initiated, and baseline to multiple follow-up changes would be evaluated.

The other design limitations relate to assessment. The results were based on analysis of self-report data. It is always possible that social desirability or recall biases affected the validity of the findings.[30] However, most studies on smoking prevention and cessation have shown that self-report data are highly valid and reliable, particularly when collected in conjunction with a biochemical measure.[19,21,22] Another limitation in assessment was the inability to differentiate smoking in and around the school setting and during school hours, from smoking in other settings and after school. Several worksite studies have suggested that smoking policy may work by delaying smoking until after hours or in other settings not affected by policy, rather than by actually changing smoking levels.[27–29,31] Although mean amount of smoking included smoking in all settings and times, it is still possible that the school policy effect is limited to reducing smoking only in and around the school.

In summary, this study suggests that smoking policy can have an effect on reducing amount of smoking by adolescents and, to a lesser extent, on smoking prevalence rates. Questions remain concerning the extent to which the effect of policy may interact with the effect of a smoking intervention in schools that are implementing both. For example, if the decreased amount of smoking in schools with a prevention-oriented smoking policy represents a conscious effort by smokers to cut down their smoking in preparation to quit, could this effort be expedited by a prevention program that is implemented in the school, or by a cessation program? In addition, little is known about the cumulative effect of multi-component smoking policies that are implemented simultaneously across several settings, e.g., a community smoking ordinance, a school smoking policy, and informal home rules about smoking.

APPENDIX A

Description of Items Analyzed for This Study

Variable	Source	No. of Items	Characteristics
Smoking			
Amount	Student questionnaire	2	No. of cigarettes smoked in last week, last 24 hours; continuous scale; high internal consistency, test-retest reliability.[19]
Prevalence Rate	Student questionnaire	3	0 = no smoking, 1 = any smoking in lifetime,† last week, last 24 hours, dichotomized from amount variables and aggregated to % of smokers in the school.
Current Staff†	Staff questionnaire	1	Continuous; estimated as % of staff currently smoking.
Observed Student‡	Staff questionnaire	1	Continuous; estimated as % of students observed smoking.
Smoking Violations‡	Clerk phone survey	1	Rated as change in last year, 1 = increased, 2 = decreased, 3 = stayed the same.
Expired Air (CO)‡	Student CO measure	1	Continuous, parts per million CO analyzed at 1 year follow-up.
Policy			
No. of Components (comprehensiveness)	Staff questionnaire	4	0 = no, 1 = yes for presence of formal (written and posted) rule about no smoking on school grounds, near school grounds, closed campus policy, formal health education plan for smoking prevention programming, summed to score 0–4; based on categories of school policy.[16,23]
Policy Emphasis	Clerk phone survey	3	1 = none, 4 = a lot for prevention, cessation, punishment.
Policy Enforcement‡	Clerk phone survey	1	1 = not at all, 4 = very regularly.
Time in Effect‡	Clerk phone survey	1	No. of years, months.
Consequences for Violation‡	Clerk phone survey	1	7 categories increasing in severity, re-scaled to 1 = detention, 2 = detention + other.
Policy on School Grounds‡	Clerk phone survey	1	0 = no, 1 = yes.
Staff Policy Awareness	Staff questionnaire	1	1 = low, 3 = high
Perceived Policy Effectiveness	Staff questionnaire	1	1 = low, 3 = high
Demographics			
Gender†	School archival records	1	% females in the school.
Race/ethnicity†	School archival records	2	5 categories re-scaled to % minority, % English as second language.
Socioeconomic Status (SES)	School archival records‡	1	% Aid to Families with Dependent Children.
	Student questionnaire	1	Father's occupation re-coded to 0 = no, 1 = professional occupation, aggregated to % in the school; high predictive validity for smoking.[20]
	Staff questionnaire‡	2	Community SES (1 = mostly professional occupation, 4 = mostly unskilled labor); summed across 2 items on school district and community SES, and standardized.
Student Achievement†	School archival records	3	% California Achievement Profile scores summed and averaged by school.
School Support	Staff questionnaire	20	1 = strongly agree, 4 = strongly disagree, summed and standardized; high validity and reliability for measuring supportive school environment.[24]

†Used for background descriptive information.
‡Used for comparison with other measures.

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Treatment of Destructive Behaviors is Focus of NIH Consensus Conference

A Consensus Development Conference on Treatment of Destructive Behaviors in Persons with Developmental Disabilities will be held in Masur Auditorium at the Clinical Center, National Institutes of Health, Bethesda, Maryland, September 11–13, 1989. This open forum will focus on the various treatments and approaches used to modify or eliminate destructive behaviors in persons with developmental disabilities. The forum will address the following questions:

- What are the nature, extent, and consequences of destructive behaviors in persons with developmental disabilities?
- What are the approaches to prevent, treat, and manage these behaviors?
- What is the evidence that these approaches, alone or in combination, eliminate or reduce destructive behaviors?
- What are the risks and benefits associated with the use of these approaches for the individual, family, and community?
- Based on the answers to the above questions, and taking into account a) the behavior; b) the diagnosis and functional level of the individual; c) possible effects on the individual, family, and community; d) the treatment setting; and e) other factors, what recommendations can be made at present regarding the use of the different approaches?
- What research is needed on approaches for preventing, treating, and managing destructive behaviors in persons with developmental disabilities?

Sponsored by the National Institute of Child Health and Human Development of NIH, the National Institute of Mental Health of the Alcohol and Drug Abuse and Mental Health Administration, the Bureau of Maternal and Child Health of the Health Resources and Services Administration, and by the NIH Office of Medical Applications of Research, this NIH Consensus Development Conference will bring together biomedical investigators, practicing physicians, consumers, and representatives of public interest groups to provide a scientific assessment of drugs, devices, and procedures and to evaluate their safety and effectiveness.

On the first two days, experts will present current scientific thinking about the diagnosis, management, and prevention of destructive behaviors in persons with developmental disabilities, and concerned voluntary organizations will be invited to make statements. On the third day, after considering the scientific evidence, the consensus panel will present its draft report and invite comments from the audience. Dr. R. Rodney Howell, Professor and Chairman, Department of Pediatrics at the University of Miami School of Medicine, will chair the panel.

To register for the conference or to obtain further information, contact:

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Containing Mental Health Treatment Costs through Program Design: A Massachusetts Study

BARBARA DICKEY, PHD, PAUL R. BINNER, PHD, STEPHEN LEFF, PHD, MARY K. UYEDA, MARK J. SCHLESINGER, PHD, AND JON E. GUDEMAN, MD

Abstract: A single site pre-post study of seriously mentally ill patients treated in a public mental health system shows that annual treatment costs can be substantially reduced with the use of day hospital treatment. Two cohorts of psychiatric patients—282 consecutive admissions to a traditional public inpatient unit in 1980, and 340 consecutive admissions to a combination of inpatient and day hospital care in 1984—were followed 12 months after admission. The substitution of the day hospital is made possible because the facility provided a dormitory residence for those who could not go home at night. Cost savings per hospital episode are about 31 per cent when the additional costs of day hospital and residence are considered. Cost shifting from inpatient to residential sites is noted, but overall mean annual costs, when all other treatment (including additional admissions), residential and family costs were included, are reduced. Readmission rates did not rise. The generalizability of the findings is limited to public mental health centers and state hospitals. (*Am J Public Health* 1989; 79:863–867.)

Introduction

After almost a decade of cost increases that have outstripped inflation, cost containment remains a primary concern of health policymakers. Cost reductions, may be gained by reducing essential services or eliminating the "fat" in the system.[1] The possibility that quality of care will suffer thus lies at the heart of the questions about the trade-offs between provision of services and cost-effectiveness in any prospective payment system.

Another approach to cost containment is to redesign the delivery of services without sacrificing their scope or access to care. This report evaluates one such effort to restructure services in a public system of mental health care and measures effect of the change on costs, access and distribution of services.

Studies have shown that less intensive care, such as day hospitalization, is at least as effective for seriously ill psychiatric patients as open-ended inpatient stays.[2–10]

With the introduction of day programs into the array of treatment modalities offered, mental health policymakers have assumed that resources would flow naturally from expensive inpatient care to less expensive, community-based care. In fact, the anticipated shift in services does not necessarily occur. Many state systems operate within a fixed budget allocation, yet for the most part they have been slow to shift resources from inpatient care to alternative services. Patterns of service delivery seem to be so entrenched that they drive the public system as much as financial payment mechanisms do. Thus, structural issues must be addressed as part of the effort to shift resources from inpatient hospitalization to alternative care.

A chance to reconsider these issues occurred when a state mental health facility was restructured to substitute day hospitalization and domiciliary care for all patients presenting for inpatient hospitalization. Earlier published reports[11,12] showed the program as both clinically effective and less costly. These early reports used cross-sectional data drawn from the institutional archives; the analyses were limited, as a consequence, to univariate statistical tests of difference, comparing different patients in treatment at several points in time. A decrease in nursing expenses was documented, but the reports did not address cost shifting. This report will address three questions:

- Can day hospitalization and other ambulatory care decrease the per patient cost of care for inpatient treatment?
- If the average cost and length of stay are reduced, what is the effect on selected aspects of care, such as the readmission rate or the use of seclusion?
- If the average annual costs per patient for intensive inpatient care are reduced through the substitution of lower cost alternatives, to what extent will costs be shifted to ambulatory and residential care? To families?

Methods

The Setting

The Massachusetts Mental Health Center is a publicly funded community mental health center serving a racially and economically mixed urban catchment area of 200,000 people. The Center provides a comprehensive range of services, including acute inpatient and long-term care, outpatient and other ambulatory treatment, and community residences. The catchment area system does not have a state hospital to which the Center can transfer patients with chronic severe mental illness. Such patients are treated at the Center, receiving the ambulatory or inpatient service that best fits their needs. The Center receives an annual allocation from the state legislature and is expected to serve (at a minimum) all the poor within the catchment area.

In fiscal year 1982, patient services at the Center were reorganized to substitute day hospitalization for inpatient care. The number of inpatient beds was reduced from 100 to 35. A 40-bed dormitory called the "Inn" was made available to day hospital patients in need of temporary housing. The amount and availability of community-based ambulatory services and residential beds remained virtually the same. The reorganization of services is described in detail

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elsewhere.[11,12] The Center has available 17.5 inpatient beds per 100,000 population, which is far below the national rate of 61.3 per 100,000 for state and county mental hospitals.[13] Residential beds within the catchment area system have remained about the same at 45 per 100,000. The area has an active aftercare and case-management program for discharged patients.

Study Design

We compared service utilization costs for one year from the index admission for two cohorts of patients treated at the Massachusetts Mental Health Center: one cohort was treated before (fiscal year 1980) and one after (fiscal year 1984) the restructuring of services and reduction of inpatient beds.

Consecutive admissions in 1980 and in 1984 included individuals sent by the court for observation and competency evaluations, admissions from civil or criminal commitments, individuals transferred from other private or public hospitals, and persons who had no catchment area assignment but were referred by local authorities. Both cohorts drawn by these admissions included patients who received day hospital treatment without any accompanying residential care, but the number of these cases was much lower in 1984 (12 as compared with 44 in 1980). We excluded from each cohort all patients transferred to another facility within 48 hours of admission or hospitalized continuously for the 12-month study period. Excluded patients accounted for 4 per cent of the 1980 cohort and 6 per cent of the 1984 cohort.

The two cohorts were alike on most sociodemographic and clinical characteristics (see Appendix), but the 1984 cohort was slightly younger, with fewer voluntary admissions, fewer high school graduates, more homeless patients, and more patients with a Major Affective Disorder. The latter is thought to reflect changes in diagnostic practice, emanating from the use of DSM III* after 1980.

Data Collection

Using chart abstracts and billing records, we recorded the type and amount of all mental health services used by each cohort member, beginning with the index admission and including all subsequent hospital and community-based services used over the next 12 months. Costs per unit of service were derived by the research staff from the general ledgers of the Massachusetts Mental Health Center and from the vendors of contracted programs. The method for deriving costs and the assumptions that underlie this work are described in an earlier report.[14] Per unit service costs derived for 1980 were inflated forward to 1984 levels using the national Consumer Price Index.

Service unit costs (e.g., inpatient per diem) were merged with utilization data at the patient level, resulting in the cost, by category of care, to treat and support each cohort member. These data were basis for calculations of per admission costs and total annual costs for each patient. This approach permits per patient costs as mean cost per user of a particular category of care and mean cost per cohort member per category of care (all cohort members did not use all services). We have chosen to report the latter because it is the only way to identify shifts of care (and costs) between categories of care.

*DSM III: Diagnostic and Statistical Manual of Mental Disorders, American Psychiatric Association.

Results

Patterns of Utilization

In the 1984 cohort, there was a shift from inpatient to day hospital treatment coupled with residence in the inn, reducing the inpatient length of stay by about half (36 days in 1980, compared with 15 days in 1984). If the "episode of care" is considered to include not just the inpatient stay but the days that followed a transfer to the day hospital, then the total stay in 1984 was 38 days.

A longitudinal analysis also revealed substitution effects. The total number inpatient days accrued by each cohort member over one year decreased from an average of 57 days (for 1.59 admissions) in the 1980 cohort to 21 days (for 1.45 admissions) in the 1984 cohort. Individuals in the 1984 cohort also spent an average of 33 days annually in the day hospital.

As a function of the substitution of less expensive care, per episode inpatient cost dropped from $6,937 to $3,397, a difference of 51 per cent; with the day hospital/inn costs added, the difference shrinks to about 31 per cent, however. If the data are aggregated on an annual level, costs for hospital care (inpatient, day hospital and, in FY84, inn) are reduced from $13,112 in 1980 to $8,204 in 1984.

In order to identify the factors contributing to the differences in annual costs, they were regressed on a model specifying age (and age-logged), race, sex, diagnosis related groups (DRGs), prior living situation, and fiscal year (the dummy variable specifying cohort). Table 1 summarizes estimates of the coefficients. The fiscal year was the major contributor to the variance. To test whether the distribution of costs affected the fit of the model, costs were logged and regressed on the same model, resulting in an R-squared of .129. Using the model as a guide in selecting variables, we tested the cost difference between cohorts after adjusting for DRG, prior living situation, and previous admissions. These cost-adjusted cohort differences shift slightly, as summarized

TABLE 1—Estimates of Variance in Annual per Patient Costs

Model Parameters	Coefficient	Standard Error
Intercept	−10409.27	23779.66
Year		
FY80	4920.96	1009.77
FY84	0.00	—
Living Situation prior to admission		
Home/self	−4124.54	4401.53
Home/family	−3360.70	4312.88
Psychiatric substance abuse facility or unit	795.29	5445.13
General hospital	835.10	5144.48
Group or nursing home	−3663.49	4532.65
Street	−4545.86	4539.34
Prison	13202.37	6855.42
Other	0.00	—
DRG		
427(P.D.)	1602.87	2113.54
427(obs)	108.16	3655.15
430(psychosis)	4814.90	1920.01
Other	0.00	—
Age	−114.79	254.48
Agelog	5927.60	9302.23
No. of previous admissions	308.02	112.99

R Squared = .114
F Ratio = 5.45
Probability = .0001

in Table 2, but support the finding of differences between cohorts.

Effect on Patient Use of Services

If those changes are detrimental to patients, they are unacceptable, regardless of the financial savings that might accrue under a restructured system. Differences, on two selected measures that might indicate patterns of care detrimental to patients, were analyzed using ANOVA to control for DRG. The results are summarized below.

Seclusion—Limited admission to the inpatient unit would restrict the census to patients with extremely difficult behavioral problems which might lead to an increase in the use of seclusion. In fact, the opposite occurred, suggesting that seclusion is more a function of administrative policy than clinical circumstances. The mean number of seclusions per patient fell from 5.13 in FY80 to 1.05 in FY84. The DRG-adjusted means differed somewhat.

Readmissions—Readmission can be calculated in two ways: over the course of 12 months, and within 90 days of discharge. First, over the course of 12 months the number of readmissions did not increase in the face of much shorter inpatient stays: 1.59 per person in FY80 compared to 1.45 in FY84. DRG-controlled means were not different. Second, a rise in the number of patients readmitted within 90 days of discharge might be interpreted as a measure of failure, whether the rapid return is a result of noncompliance with discharge plans, inadequate discharge planning, inadequate hospital treatment, premature discharge, or reoccurrence of the disorder despite every effort on the part of the clinicians and the patient. We found that from 1980 to 1984, the rate of readmissions within 90 days of discharge, calculated as the proportion of all admissions, did not change: 13 per cent of all readmission in FY80 and 14 per cent in FY84 occurred within 90 days of discharge.

Discharge Status—Data collected on the patients' discharge disposition after the last admission during the study period showed little change across the two cohorts, as summarized in Table 3, with about half the discharges to home and about 11 per cent to community residences. From 1980 to 1984, however, there were small increases in the number of discharges to court and the number of patients who left against medical advice, escaped, or eloped.

Cost Shifting: Annual Community-Based Service Use

Expenditures generated by the two cohorts through the use of mental health services, and residential care during the study period are summarized in Table 4.

Inpatient Costs—Although the per unit cost per day of care increased slightly, the average annual cost per cohort member was reduced because utilization changed: both the length of stay and the number of admissions per patient dropped.

TABLE 2—Adjusted Cost Differences between FY80 and FY84

Adjustment	Point Estimate	95% Confidence Interval
Living situation prior to admission	$4,965	$1,676–$8,254
DRG	$5,007	$2,043–$7,971
No. of previous admissions	$4,703	$4,511–$4,895

TABLE 3—Discharge Disposition

Disposition	FY80*	FY84*
Living at home	44%	45%
Nursing home	4%	1%
Psychiatric unit or freestanding psychiatric hospital or not discharged from MMHC	14%	9%
Detox or substance abuse unit	—	<1%
Med-surgery, general hospital	2%	3%
Group Residence, supervised apartment	11%	10%
Court	4%	8%
Prison, forensic unit	<1%	2%
Shelter or street	3%	<1%
Escape, elopement	8%	12%
Left state	7%	5%
Died	<1%	<1%

*Sum to less than 100% due to rounding error.

TABLE 4—Average Annual Costs per Patient for Mental Health and Related Services, 1980 and 1984

Costs	1980 (n = 282)	1984 (n = 340)	% Change
Inpatient	$11,028	$4,892	−56
Day hospital	2,084	2,151	+3
Other ambulatory	117	110	−7
Residential	1,113	2,228*	+100
Forensic	346	165	−52
Family and social burden	478	204	−57
Total	$14,866	$9,750	−34

*Includes The Inn

Day Hospital Costs—Although the mean costs per cohort member appear to be similar, they represent an example of changes in both per unit costs and in utilization: per unit costs in the 1980 day hospital were much higher, for several reasons: the census was routinely below capacity, the overhead and capital costs were relatively high, and the staffing configuration included more costly professionals. In 1984, the per unit costs were lower because overhead and capital costs were reduced along with the level of professional staffing. At the same time, census was at capacity. Thus, cost shifting in this category of care does not appear to be evident.

Other Ambulatory Costs—These costs account for the use of the Center's ambulatory daytime programs, such as prevocational and rehabilitative services, and social activities. Availability of providers, methods of reimbursement, and operation of the outpatient department and per unit cost of service did not change between the two studies periods. Utilization remained about the same.

Residential Costs—The costs in 1984 rose, as expected, to more than double the 1980 figure. Cost shifting from the inpatient unit to this category of care is evident, although the increase was due not just to the use of the inn. The per unit inflation-adjusted costs of residential housing increased substantially over the period because of rising property values in Boston and improved standards of housing for the mentally ill.

Forensic Hospital Costs—The state forensic facility for the criminally insane is managed by the Department of Corrections but provides a minimal level of treatment for

mental illness. It does not constitute a back door for long-term difficult patients as transfers occur only when that facility agrees to the admission. Mean costs dropped from 1980 to 1984 because of a decrease in utilization.

Family and Social Burden—Costs were calculated as a surcharge to each day hospital or daytime ambulatory treatment day for which no accompanying inn or residential community night were identified. For patients who lived at home during treatment, we estimated the cost to the family based on the work of Frank,[15] who used both personal expenditures and time spent on behalf of the ill family member to arrive at a daily cost to the family. In some instances, patients were living on their own or in a shelter. Costs for these nights were derived from shelter budgets or from SSI/SSDI (supplemental security income) transfer payments, as estimates of minimum costs.

Discussion

Our study has demonstrated a substantial reduction in inpatient costs within a public system of mental health care where both the organizational and fiscal structure created pressures for efficiency. The results suggest that costs can be reduced by differentiating levels of care within a service system, rather than by relying solely on the arbitrary manipulation of reimbursement mechanisms. The study also demonstrated expected shifts in service use from inpatient care to residential programming and day hospital care, reducing somewhat the savings that accrued the reduction in inpatient care.

Our study only begins to address the many complicated questions that arise when attempting to measure the effects of structural change within a system. Findings from a single-site study are undoubtedly compromised by unmeasured site-specific variance. Despite its methodological limitations, our findings confirm what has been the hypothesized cost outcome of the use of hospital "alternatives." Where other studies have selected certain special populations for care in alternate settings, this study took advantage of a program that made no exceptions in program admission, and included all patients needing inpatient care. It remains to be tested whether or not day hospitalization can substitute with equivalent results for individuals treated in other systems.

Certain structural problems will need resolution before widespread use of day hospitals occur. First, centralized administrative and clinical accountability across local catchment area services are thought to be a necessary (but not sufficient) characteristic of effective public mental health systems.[16] Most local systems lack this accountability and are instead a collection of quasi-independent agencies that establish their own program goals for a particular population of clients, and maintain control of admission and discharge from their own programs. Active case-management programs may reduce the access barriers to these programs, but smooth flow of services to the chronically mentally ill remains a major problem in this country.

Second, in many states, a significant problem exists in the management of human resources when phasing down labor-intensive state hospitals and reallocating staff to different programs. The problems arise for several reasons: the hospitals are often in rural isolated communities where the state is the major employer, or there may be strong union activity supporting continuation of jobs historically within the scope of contracts negotiated with the state.

Our findings may be important in considering special financial arrangements for treating the severely and chronically mentally ill in the public sector. One such arrangement that has been proposed is a capitated mental health organization offering a continuum of comprehensive services for a fixed per capita payment.[16–18] The fixed payment provides an incentive to move clients to less costly services as quickly as possible, increasing service efficiency.

Although public capitated systems of care for the chronically mentally ill are rare, one has just been started in Rochester, NY* and several others are in planning stages.** Data from these systems are not yet available. In the absence of evidence from these plans of the likely effect of incentives to provide alternatives to hospitalization, the findings from this study might be of some value. For example, the shortage of public psychiatrists and psychiatric nurses might be alleviated by a system that made the most of their special training, in inpatient units. Other mental health professionals, with limited use of psychiatrists and nurses, could be the primary work force in less medically oriented treatment settings.

In summary, we believe that the key to reducing the use of inpatient treatment and encouraging the use of alternative services lies in structural changes, not in fiscal incentives, that target services more efficiently by matching clinical need with service delivery, rather than by reducing by some fraction all services to all patients. The latter does not address the current situation in the public sector where some patients receive more intensive services than they need and others need more than they are currently receiving.

Second, local service systems might be more efficient if they were clinically and administratively centralized.[19] This combination of factors, rather than fiscal incentives alone, can be used to reshape clinical delivery systems to ensure that the seriously mentally ill are adequately and efficiently served under acceptably standards of care.

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APPENDIX
Characteristics of Patients in the 1980 and 1984 Cohorts

Patient Characteristics	FY1980 (N = 282)	FY1984 (N = 354)
	%	%
Sex		
Male	51	57
Age		
Mean (years)	41 years	35 years
Race		
White	72	65
Black	19	23
Hispanic	6	9
Asian	1	3
Other	1	0
Legal		
Involuntary	33	44
Education		
<High school	28	36
Completed HS	26	22
Trade/technical	5	5
Some college	22	21
Undergraduate degree	13	10
Some graduate	3	2
Graduate degree	3	5
Marital Status		
Not married	90	84
Primary Diagnosis		
Schizophrenia	37	34
Affective disorder	22	32
Schizoaffective	8	7
Substance abuse	3	2
Mental retardation	—	1
Personality disorder	22	19
Organic brain syndrome	4	1
Other	6	3
No. of Previous Psychiatric Admissions		
0	19	25
1	13	16
2–5	37	36
6–10	25	15
11–15	5	5
>16	2	4
Most Recent Living Situation		
Residence (self)	26	23
Residence (family, friend)	49	45
Psychiatric or substance abuse unit	4	1
General hospital	—	5
Community residence, nursing home	13	10
Street or shelter	7	14
Prison	1	1
Other	1	1

Different Views

The President's Commission on AIDS: What Did It Do?

KRISTINE M. GEBBIE

Introduction

It can be hard to comment on something you have experienced intimately as one of the 13 members of the Presidential Commission on the HIV Epidemic.

The distance of several months is helpful, however, and it has been educational for me to reflect on the process.

The Commission was made up of 13 individuals. By non-mutually exclusive categories, we were eight men, five women, one Black, one gay, five physicians, two nurses, two academics, one minister, one legislator, one publisher, five business executives, one government official, one sailor; geographically, seven from New York/Washington, three from the Midwest, two from the West Coast, one Southerner. There is no way we could be representative of the nation, but we tried to be conscious of all segments of our society. Several were entities "known" to the Reagan Administration, some not; some had congressional support, some were known to their colleagues, some were a category about which it could be said "we've got to have one of them," etc. What I brought to the Commission's process was exposure to the public health system, separation of illness care from prevention, and perhaps some broadening of awareness of the complexity of experience with human immunodeficiency virus (HIV) and public health policy within state health officials association, and in my own State. Perhaps it is important that I did *not* come from New York, California, Florida, New Jersey, Texas. Many people do not identify with those states, and although they have led in this epidemic (both in problem and solutions) their experience must be translated to be useful to others.

The charge to the Commission was enormous: "advise the President and Secretary of Health and Human Service and other relevant cabinet heads on public health dangers including medical, legal, ethical, social and economic impact, from the spread of the HIV and resulting illnesses."[1] Specific areas requested included: "recommend measures that Federal, State and local officials can take to protect the public . . . to assist in finding a cure and . . . to care for those who have the disease."[1] The Commission was further charged to "evaluate efforts by educational institutions . . . analyze efforts currently underway by Federal, State, local authorities, examine long term impact on care system . . . review U.S. history of dealing with communicable disease . . . evaluate research activities . . . identify future areas of research . . . examine policies for development of drugs . . . assess progression of AIDS among the general population and subgroups . . . study legal and ethical issues . . . review the role of the US in the international pandemic."[1]

The Chairman, when I was appointed, was an Admiral asked by his Commander-in-Chief to complete a mission. There were two key elements on which he insisted: 1) credible testimony in public hearings would be the basis for recommendations (rather than any preconceived ideologies of Commissioners); and 2) the regular appearance of HIV-infected persons would put more abstract issues into the human perspective.

Things the Commission did not do

It did not do much, if anything, to bring the President into the HIV response, possibly because he is so shielded by the conservative voice. To no avail, we criticized the Reagan Administration, suggested attitudes and approaches not comfortable to its leaders, asked for monies that had not been authorized. It was probably too much to hope that there would be a sudden turn around. Yet it might have been a nice finish if we at least could have met in the Rose Garden.

Neither did we break the Congressional log jam or budget. There was no immediate action to pass the Americans with Disabilities Act[2] or pour monies into drug treatment. Yet the Commission was clear that both actions were critical first steps. A word about each:

- The Americans with Disabilities Act recommendation was the cause of the most serious disagreement in the final report, with a split vote. I believe the debate was not about fighting or not fighting discrimination, but on this particular act as the vehicle. It takes a broad, "civil rights" approach to *all* disabilities, not setting the HIV-infected apart. The majority of the Commissioners viewed it as the most equitable and efficacious way to proceed. Chief opposition comes from the business community which saw this approach as "too costly."
- Massive increase in drug treatment (to the level sometimes called "on demand") requires re-thinking substance abuse as a chronic condition which may need treatments at several points in an individual's life. We must become serious about attacking addictions with *all* available weapons, not just police action and "just say no" campaigns.

We did not really resolve what I think is going to be a major subject of debate in the coming months: the appropriate response to a health care worker exposed to a potentially infectious substance, and the appropriate behavior of an infected health care worker. The Commission report includes some nice paragraphs on the subject, talking about "workers' right to know" (long a favorite public health phrase), confidentiality, prospective policies regarding infected workers. Listening to physicians, nurses, EMT's (emergency medical technicians), their spouses, lovers, employers over the past few weeks, it becomes clear to me that this issue is far from

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Public Health Briefs

Sero Survey of Human Immunodeficiency Virus Infection in Women at a Family Planning Clinic: Absence of Infection in an Indigent Population in San Francisco

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Abstract: To determine the seroprevalence of human immunodeficiency virus (HIV) in contracepting women, 1,308 consecutive attendees at an inner-city family planning clinic completed a risk factor questionnaire and were then asked to have an HIV test. None of the 1,000 women tested was seropositive. Preventive efforts in family planning clinics are likely to be productive because few contraceptors, including those with risk factors, are already infected. The factors which determine differences in HIV positivity between childbearing and contracepting women, who are similar in demographic and HIV risk characteristics, warrant further investigation. (*Am J Public Health* 1989; 79:883–885.)

Introduction

Concern is expressed[1–3] over the possibility that the human immunodeficiency virus (HIV) will rapidly become more prevalent in women who can then pass it to their sexual partners[4–6] and babies.[7–11] Rates of HIV positivity in childbearing women range from less than one per 1,000 in suburban and rural populations[12] to as high as two per 100 at inner-city municipal hospitals.[13] Such determinations of HIV infection in women have generally relied on cord blood or heel stick surveys. Few investigations have screened actively contracepting women and avoided, or accounted for, possible selection bias. None have systematically assessed the prevalence of risk factors in such a population. Since 55 percent of US women aged 15 to 44 are active contraceptors,[15] a family planning clinic population may more accurately reflect the HIV status of women of childbearing age than do other populations. This seroprevalence study assessed the level of HIV infection in relation to risk factors for this infection at a family planning clinic serving the indigent women of a city where HIV infection is endemic in homosexual males.

Methods

The study was conducted from November 1987 to April 1988 in an urban hospital-based family planning clinic. In 1987, 93 percent of the clinic's 3,116 active patients lived in households with incomes beneath the "poverty level"; 12 percent used condoms with or without spermicidal foam, 20 percent used diaphragms and other barrier methods, 65 percent used oral contraceptive pills, and 3 percent used IUDs (data available on request to authors). As part of routine evaluation, all women attending the clinic completed a confidential, self-administered questionnaire, had blood drawn by lab personnel, and were then seen by a clinician. The questionnaire was designed to elicit a history of intravenous (IV) drug use, blood transfusions between 1979 and 1985, artificial insemination between 1979 and 1985, sexual practices including sexual contact with persons diagnosed with AIDS (acquired immunodeficiency syndrome), ARC (AIDS-Related Complex), or hemophilia, sexual contact with IV drug users, sexual contact with bisexual men, and sexual contact with more than five partners in any one year since 1979. Clinicians reviewed the women's responses with them and provided information on HIV infection, prevention, and implications of antibody testing. The women were then asked to permit an HIV test of their blood. The study was approved by the Committee on Human Research of the University of California, San Francisco.

If women consented to an HIV test, an anonymous number coding system permitted them to obtain test results. Blood tests were performed only for women who consented, but demographic and risk factor information was collected for both the women who agreed to testing (participants) as well as those who did not (nonparticipants). Blood specimens were collected aseptically and stored at 4°C pending transport for analysis.

The Viral and Rickettsial Disease Laboratory, Division of Laboratories, California State Department of Health Services analyzed the blood samples using EIA Dupont and Genetic Systems kits. The laboratory analyzed repeat positives with IFA reagents, and if EIA and IFA did not agree, performed Western Blot.

Results

A total of 1,308 women completed questionnaires during the study period, which continued until 1,000 women agreed to participate in the study. There were no significant differences in demographic characteristics of the 1,000 participants and the 308 nonparticipants (Table 1). Participants were more likely to report histories of STDs (sexually transmitted diseases) (Table 1), riskier behavior, and have at least one AIDS risk factor than were nonparticipants (Table 2). Although 14 percent of the participants claimed one or more risk factors, none had HIV antibodies.

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TABLE 1—Serosurvey Participants and Nonparticipants by Selected Demographics

Characteristics	Participants (N = 1000)	Nonparticipants (N = 308)
Ethnicity		
% Latino	59.3	56.8
% White	17.3	17.2
% Black	15.3	15.9
% Asian	5.9	6.8
% Other	2.2	3.2
Mean Age (years)	26.3	26.4
Mean Live Births	1.4	1.5
Mean Elected Abortions	0.9	0.8
History of STDs (%)	13.7	6.5

TABLE 2—Number of Risks and Specific Risks Reported by Serosurvey Participants and Nonparticipants

	Participants (N = 1000)		Nonparticipants (N = 308)	
Number of Risk Factors	N	%	N	%
One Risk Factor	101	10.1	15	4.9
Two Risk Factors	30	3.0	3	1.0
Three Risk Factors	7	0.7	0	0.0
At Least One Risk Factor	138	13.8	18	5.8
No Risk Factors	830	83.0	279	90.6
Some Not Reported; all others negative	32	3.2	11	3.6
Risk Factors				
IV Drug Use by Self or Partner	26	2.6	0	0.0
5 or More Partners/Year	72	7.2	7	2.3
Partner with AIDS or ARC	1	0.1	0	0.0
Blood Transfusion, 1979–85	19	1.9	3	1.0
Bisexual Partner	11	1.1	0	0.0
Partner Possibly Bisexual	53	5.3	11	3.6

Caucasians and Black women had much higher incidence of risks than did Latinos and Asians (30.1 percent and 20.3 percent versus 7.9 percent and 6.8 percent, respectively) (Table 3). Caucasians were the most likely and Latinos least likely to report multiple AIDS risk factors. A total of 237 women (24 percent) chose to return for their results. This choice was positively correlated with the level of risk: those with greater numbers of risks, those with self or partner IV drug use, and those who had received blood transfusions were the most likely to return for the HIV test results.

Discussion

There are several possible explanations of the absence of seropositives in our study population: We surveyed a low-risk population; seropositivity is uncommon among San Francisco women in general, including those who report risk factors; or women attending family planning clinics may behave in ways which place them at lower risk than their counterparts in the childbearing population.

Overall, fewer women reported risk factors in this study than in other studies. The fact that the present questionnaire was self-administered may have further increased the chance that some risks were not reported, due to confusion, oversight, or embarrassment. However, other sero surveys have suggested that self-reported risk factors are unreliable predictors of HIV positivity,[3,11,13] so this risk assessment of San Francisco women may not represent their actual risk. Because 85 percent of the women attending this clinic used contraceptive methods not known to protect against HIV infection (pills, IUD, or diaphragms), the use of contraception by itself seems an unlikely explanation of lower risk. Only the history of intravenous drug use was a significant predictor in Quinn, *et al*'s,[3] multivariate regression analysis of risk factors. The fact that most IV drug users do not use traditional family planning services means that we did not survey many of the women who are at the greatest risk.

These observations suggest that HIV seropositivity is uncommon in San Francisco family planning clinics, and thus universal HIV screening in these clinics is unwarranted. Instead, these are ideal places for educational efforts to reduce heterosexual transmission of the virus. These results also suggest that further examination of the behavior characteristics which differentiate women who attend family planning clinics from those in the childbearing population might help identify patterns of behavior which place women at lower risk of heterosexual HIV transmission. Because previous studies[3,12,13] were conducted in East Coast cities only, studies of other West Coast cities are needed to determine if HIV infection differs between contraceptors and the remaining female population of childbearing age. In addition, innovative outreach programs are urgently needed to educate those women who use, or whose partners use, IV drugs of the risks of HIV infection to themselves and their fetuses, and to provide contraception to such women.

ACKNOWLEDGMENTS

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TABLE 3—Risk Factors of HIV Infection by Ethnic Group

Risk Factors	Latino (N = 593) %	White (N = 173) %	Black (N = 153) %	Asian (N = 59) %	Other (N = 22) %	Total (N = 1000) %
IV Drug Use by Self or Partner	0.7	8.7	3.9	1.7	0.0	2.6
5 or More Partners/Year	2.9	19.1	12.4	3.4	4.5	7.2
Partner with AIDS or ARC	0.2	0.0	0.0	0.0	0.0	0.1
Blood Transfusion, 1979–85	1.9	3.5	0.0	3.4	0.0	1.9
Bisexual Partner	0.5	3.5	1.3	0.0	0.0	1.1
Partner Possibly Bisexual	3.5	9.8	7.2	1.7	13.6	5.3
At Least One Risk Factor	7.9	30.1	20.3	6.8	18.2	13.8
Two Risk Factors	1.3	8.7	3.3	3.4	0.0	3.0
Three or More Risks	0.2	2.9	0.7	0.0	0.0	0.7

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Stress Reduction Training Changed Number of Sexual Partners but Not Immune Function in Men with HIV

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Abstract: We tested the impact of stress management training on sexual behavior and immune functioning in 64 gay men infected with human immunodeficiency virus (HIV). Subjects randomized to the stress management group met for eight two-hour sessions and one all day retreat to learn systematic relaxation, health behavior change, and stress management skills. Compared to those randomized to a wait list control, treatment subjects reported significantly fewer sexual partners in the prior month at post-test (1.10 vs 2.29 for controls). There were no differences between groups in lymphocyte numbers and function. (*Am J Public Health* 1989; 79:885–887).

Introduction

Significant associations between stressors, decrements in immunological function and clinical disease have been reported.[1,2] Relaxation techniques,[3] humorous film stimuli,[4] and operant conditioning procedures[5] may contribute to immunoenhancement. Levy and Ziegler[6] proposed that HIV (human immunodeficiency virus) infection is an opportunistic disease, progressing to clinical AIDS (acquired immunodeficiency syndrome) only in those who are immunocompromised by other factors. To date, several co-factors for disease progression have been identified, including time of infection, age, incidence of gonorrhea, and use of hallucinogens.[7–9] Antigen exposure, drugs and alcohol use, nutrition, and health habits as well as psychosocial stress are also posited cofactors.[2] Tension reduction has been related to levels of sexual activity which may transmit HIV.[10] This study was designed to test experimentally the hypothesis that group stress reduction training will lead to improvements in immune function and reductions in unsafe sex.

Methods

One hundred twenty men responded to a brief article placed in a local San Francisco gay newspaper asking for volunteers. Subjects were selected who: were seropositive for HIV; had not experienced oral candidiasis, hairy leukoplakia, recurrent herpes infection, or herpes zoster; and who were not already practicing meditation regularly. A total of 64 men (mean age = 34.9 years) who met the criteria completed baseline questionnaires and venipuncture. Subjects were then randomized into treatment or control groups. All subjects signed a statement of informed consent approved by the Human Subjects Review Committee of the University of California-San Francisco. Immediately following the termination of each treatment, all experimental and control subjects were re-assessed on the behavioral and immunological measures. Immunological assays included enumeration of lymphocytes and subsets.[11] In addition, we completed a number of functional measures of immunity. These included: Natural killer cell function,[12] lymphocyte response to Concanavalin A (ConA) and Candida antigen[13] and to cytomegalovirus (CMV),[14] and serum Immunoglobulin A (IgA).[15] Laboratory staff were blinded as to subject name and assignment.

Stress reduction training was completed in two-hour sessions once per week for eight weeks, with one all day retreat after week 4. The goals of the sessions were:

- learning systematic relaxation—through use of group time and take-home tapes, to attain some systematic relaxation for each individual every day;
- health habit change—including execution of behavioral contracts around diet, rest, exercise, drug and alcohol use, and smoking;

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- learning skills for managing stress—including the stress of being seropositive through the use of group discussion, teaching, and modeling.

We used analysis of covariance with the baseline as the covariate to test post-treatment differences.

Results

Presented in Table 1 are post-test differences in adjusted means with 95% confidence intervals. Treatment subjects reported fewer sexual partners at post-test than did control subjects. Only 14 per cent of subjects reported practicing unsafe sex at baseline, with no differences between groups. At post-test, 15.9 percent of subjects reported practicing unsafe sex, again with no difference between groups. There were no post-test differences in measures of immune number or function.

Discussion

Reducing the number of sexual partners may be important for HIV seropositives. Polk, *et al*,[16] reported that number of partners was a risk factor for disease progression among infected men. Hessol, *et al*,[9] reported that incidence of gonorrhea predicted disease progression. The ability of this kind of treatment to reduce multiple partner sex and unsafe sex should be explored further.

There are several reasons why these negative immune outcomes may have occurred:

- We measured immune function one to four days following the termination of the group. Sadness at termination may have caused deterioration in immune function.
- Immune function may not be influenced by stress or stress reduction in the presence of HIV infection.
- Single measures of immune function may not be sensitive to the subtle effects of stress management. Multiple assessments may be needed to get stable measures.
- The psychosocial intervention may not have been sufficiently potent. Extremely intense interventions are needed to bring about clinically significant changes in risk factors for ischemic heart disease.[17]
- The nature of the group may also have induced stress as the leaders attempted to reduce it.

Whereas the presence of other infected men helped individuals to recognize and confront the psychological issues related to AIDS, participation also focused closely on immune function, which can provoke anxiety. Awareness of HIV seropositivity, a screening variable for the study, may further increase anxiety. This may be indicative of the dilemma faced by members of the San Francisco community: as individuals seek to cope with an epidemic in their environment and protect their own health, the anxiety associated with coping collides with the need to reduce stress and conduct good health maintenance.

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TABLE 1—Pre- and Post-Treatment Sexual Behavior and Immune Function for Experimental (E) and Control (C) Subjects

Variables		Pre-Treatment		Post-Treatment		Difference of Adjusted Post-Treatment Means	95% CI
		Mean	(SD)	Mean	(SD)		
Number of Sexual Partners	E	1.41	(1.80)	0.50	(0.50)	1.20	(.14, 2.28)
	C	1.09	(2.44)	2.29	(3.79)		
Absolute number of T-Helper Cells per cu. mm blood	E	460	(195)	447	(188)	46.94	(−85.31, 51.86)
	C	570	(170)	523	(214)		
Helper to Suppressor Ratio	E	0.69	(0.34)	0.63	(0.32)	0.05	(−.022, .[illegible])
	C	0.75	(0.29)	0.72	(0.31)		
NK Assay (% cy Fo Toxicity)	E	34.0	(15.0)	41.4	(15.5)	2.87	(−9.67, 3.92)
	C	34.0	(13.0)	38.5	(11.5)		
Mitogen Response	E	21351	(12240)	13856	(9114)	1074.99	(−2850.32, 5000.30)
	C	24124	(13616)	16194	(10292)		
CMV Antigen Response	E	38.0	(113.3)	16.8	(29.3)	1.54	(−11.18, 14.25)
	C	31.6	(47.6)	21.1	(37.8)		
Candida Antigen Response	E	13.8	(29.4)	16.9	(24.3)	0.31	(−10.48, 11.10)
	C	11.0	(15.2)	16.1	(21.2)		
IgA (Mg/dl)	E	248.6	(130.3)	262.4	(433.0)	1.09	(−15.68, 17.86)
	C	210.5	(86.8)	224.7	(100.2)		

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Characteristics of Self-Selected Responders to a Health Risk Appraisal: Generalizability of Corporate Health Assessments

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Abstract: Selected characteristics and total medical claims of health risk appraisal (HRA) responders and non-responders were compared in a sample of employees having a three-year employment and claims history. HRA responders were younger and more likely to file medical claims than non-responders. Although mean medical claims were greater for HRA responders than non-responders, when adjusted for age and sex this difference reflected the proportion of employees reporting claims, not a difference in the claims amount. (*Am J Public Health* 1989; 79:887–888.)

Introduction

There is conflicting information about the characteristics that influence participation in health promotion activities,[1] although it has been suggested that participants are healthier than non-participants.[1–3] This assumption has not been documented by differences in health status or health costs however. To address this concern, the present investigation tests the null hypothesis that independent of their age, sex, or job type, health risk appraisal (HRA) responders are similar to HRA non-responders with respect to health costs.

Methods

The generalizability of HRA data in our present setting was determined by the degree of similarity between the self-selected HRA responders and non-responders.

The subject population consisted of full-time employees of the Travelers Companies home office, having at least a three-year employment history as well as group health insurance (not health maintenance organization) (N = 6,989). All eligible employees were requested to complete a voluntary HRA instrument, the Lifescore Plus questionnaire. Comparisons between HRA responder (2,600) and HRA non-responder (4,389) group characteristics were based on personnel information, sex, age, and job type (salaried or hourly).

The average amount paid for medical claims over the three-year period prior to administration of the HRA was calculated for each insured employee. Only claims for the subjects (not for spouses or dependents) were included. The group insurance plan covered almost all medical expenses (with limits on out-of-pocket co-payments) including ambulatory care, medications, surgery, and hospital and home health care services, up to a lifetime maximum of $2 million. In addition, the insurance plan paid 80 per cent of inpatient psychiatric expenses and 50 per cent of outpatient mental health expenses. Employees were required to pay $100 cash deductible per year and additional deductible amounts for each hospital confinement. Expenses that were not covered included routine medical tests, service-related disability expenses for veterans, and expenses covered under workers' compensation. Insured employees having no record of filed claims were assigned an average of zero dollars.

Because the claims data were extremely skewed, several forms of this variable were examined. In one form, to account for the disproportionate weight carried by the highest cases, the top 1 per cent of the cases (ranging from $5,971 to $41,711) were removed (N = 70). In another form, average claims values were adjusted (cube root transformation) to reduce the extremely skewed nature of the data. Mean claims for the HRA responder and HRA non-responder groups, in their adjusted and unadjusted forms, were compared using a hierarchical multiple regression to remove the influence of age and sex differences.

Results

HRA responders differed from non-responders in several respects. HRA responders (39.7 ± .21 years) were younger on average than non-responders (41.1 ± .17 years). A greater proportion of salaried and a lesser proportion of hourly employees responded to the HRA (61 ± 0.95 per cent salaried in the HRA responder group vs 48 ± 0.75 per cent). Males and females were equally represented in both groups. Interestingly, HRA responders were more likely to file health

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claims (78 ± 0.8 per cent) than non-responders (72 ± 0.7 per cent). Excluding the outlying cases, mean claims for the HRA responder and non-responder groups were $424.52 (median = $89) and $399.67 (median = $77), respectively.

Overall, HRA responders experienced higher average claims than non-responders after means were adjusted for age and sex differences. Type of employment (salaried or hourly) had no independent bearing on claims costs. Table 1 shows the estimated cost difference between HRA responders and non-responders for all insured subjects using claim costs both in their original form and without the outlying 1 per cent of the cases. Using the more conservative estimate, the HRA responder is expected to cost almost $40 more per year than a non-responder of the same age and sex. However, when only those filing claims were compared the difference in average claim amounts was no longer evident.

Discussion

The results of these simple demographic and medical cost comparisons contrast with previous reports concerning participants in worksite health promotion activities. First, there were distinct characteristic differences between those who selected to respond to HRA and those who selected not to respond. Whether these differences are consistent with HRA participation in other settings is not known.

Second, earlier suggestions that participants in organized health promotion programs are healthier than non-participants and may use fewer medical services were not substantiated in our research population. HRA responders were more likely than non-responders to file medical claims, yet for those who filed claims, costs were similar. Put simply, the health profile for HRA responders cannot be generalized safely to the larger employee population.

TABLE 1—Estimated Medical Claims Differences between HRA Responders and Non-Responders (Holding Age and Sex Constant)

	Cost Difference (Mean Responder Claims—Mean Non-Responder Claims)	
	Cost Difference	(95% CI)
All insured (including outliers)	$70.55**	($35.34, $105.76)
All insured (excluding outliers)	$38.43**	($19.11, $57.75)
Those filing claims (including outliers)	$36.81	($−9.24, $82.86)
Those filing claims (excluding outliers)	$ 3.87	($−20.89, $28.63)

**Cost differences significantly greater than zero.

The observed relationship between HRA response and health claims lends itself to several possible explanations. Perhaps HRA responders are overly concerned about health and more likely to seek health services. On the other hand, those in poor health may have greater motivation to respond to HRA. Another possibility might be that, ignoring health status, some individuals dislike paperwork and are less likely to fill out forms of any kind, HRA instruments and reimbursement forms included. Whatever the explanation, HRA responders seem to behave differently than non-responders.

Results of the comparisons based on transformed claims data were consistent with the findings above, with respect to the hypotheses that were tested. Because the data conform to the assumptions on which linear regression are based, the transformed claims produced better model fit and reduced the inflation in estimates of claims differences. These issues were not addressed here, but pertinent findings are available upon request to the authors.

We recognize that the claims records used in this investigation represented only those services that were reported by the employees, and that specific diagnoses could not be identified. Almost all medical expenses were covered by the plan, with a few exceptions that had no obvious relevance to the comparisons made in this investigation. Due to normal procedures, a claim must have been filed by the employee in order for the employee to receive payment for expenses. We cannot know whether true health expenses were underestimated because of claims not filed.

Because it may not be possible to generalize to HRA non-responders, health promotion practitioners should place a high priority on encouraging high response rates from their employees. In addition, repeated administration of HRA within the same company may reflect both health status changes and changes in self-selection so that a true indication of short-term changes in behavior can be inferred only by comparing individuals with themselves.

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Personality Risk Factors for Cocaine Abuse

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Abstract: To evaluate the role of personality in cocaine abuse, 59 adults meeting DSM-III criteria for cocaine abuse were compared to similar-aged non-cocaine alcohol abusers and community controls on a DSM-III measure of personality. Cocaine abusers were more likely than non-cocaine alcohol abusers to display narcissistic personality traits (Odds ratio 6.86, 95% CI = 4.52, 15.60). (*Am J Public Health* 1989; 79:891–892.)

Introduction

Although recent surveys suggest a decrease in recent cocaine use among high school seniors,[1] cocaine abuse continues to be a current problem for over one million Americans.[2] With widespread cocaine abuse, cocaine-related emergencies and death have tripled over a five-year period.[3] The illusion of cocaine as a "safe" drug has evaporated as pure, more potent, and less expensive forms of cocaine have evolved. Cocaine abuse is now recognized as a significant cause of psychiatric and medical morbidity.[4]

Psychiatric epidemiologic surveys point to common associations (comorbidity) between major psychiatric disorders and substance abuse.[5] For alcohol abuse in the community, antisocial personality disorder is a frequent comorbid condition.[6] The role of personality in cocaine abuse is less well-defined. Uncontrolled descriptive studies have found high rates of some personality disorders in clinical samples of cocaine abusers.[7] To further evaluate personality as a risk factor for cocaine abuse, we studied the personality of a group of cocaine abusers compared to non-cocaine alcohol abusers and community controls.

Methods

Cases and psychiatric controls were obtained from an unselected series of 229 substance abusers seeking inpatient treatment at one of three Iowa City, Iowa treatment facilities. Cocaine abuse was defined according to the criteria of the Diagnostic and Statistical Manual of Mental Disorders, Third edition (DSM-III), using the Diagnostic Interview Schedule.[9] Fifty-nine (32.9 per cent) of the sample met criteria for cocaine abuse and were identified as cases. A group of 47 similar-aged substance abuse controls were identified by meeting DSM-III alcohol abuse or alcohol dependence criteria. A minority of substance abuse controls had used cocaine but none met criteria for abuse.

The Personality Diagnostic Questionnaire (PDQ) was administered to all cases and controls.[10] The PDQ assesses personality using 11 non-exclusive categories in a 152-item forced choice, true-false self-report questionnaire. The DSM-III describes 11 personality categories grouped into three clusters based on similar psychological constructs with cluster A denoting eccentric features, cluster B denoting flamboyant features, and cluster C denoting anxious features. Although the methodology of assessing personality by DSM-III criteria has only recently emerged, the PDQ has some evidence for test-retest reliability for many personality disorders[11] and has demonstrated prognostic validity in patients with major depression.[12] Additionally, the PDQ has been used in psychiatric clinical populations[13] and has been used in a community survey.[14] Sixty-nine similar-aged community controls without cocaine abuse were identified from this previous community survey using the PDQ. Cases and both control individuals were all identified blind to their PDQ results.

The distribution of personality disorder and traits is influenced by the respondent's gender.[15,16] The Mantel-Haenszel chi-square statistic was calculated to correct for group sex differences. Odds ratios were calculated with 95% confidence intervals using the method of Woolf[17] with calculations of combined sex-weighted odds ratios.[18]

Results

The count and ages (standard deviation) for individual groups by sex were:

cocaine abusers—men N = 36, 28.0 (6.3 years), women N = 23, 26.0 (6.4 years)

alcohol abusers—men N = 22, 31.0 (6.7 years), women N = 25, 28.6 (6.2 years)

community controls—men N = 30, 29.8 (4.5) and women N = 39, 22.2 (4.3).

The three groups were all largely Caucasian in race: cocaine abusers 84.7 per cent, alcohol abusers 87.2 per cent, and community controls 94.2 per cent.

Personality comparisons between groups are shown in Table 1. Narcissistic personality was more common in cocaine abusers compared to alcohol abuse controls (32.2 per cent vs 6.4 per cent, $X^2 = 10.92$, Odds ratio = 6.98, 95 per cent CI = 1.26, 37.1). Additionally, cocaine abusers were more likely than alcohol abusers to report schizoid features (13.5 per cent vs 0 per cent, Fishers exact one-tailed p = .042). Cocaine abusers demonstrated multiple personality features compared to community controls. Cluster B disorders were especially more prevalent among the cocaine abusers compared to community controls (79.6 per cent vs 10.11 per cent, $X^2 = 61.4$, Odds ratio = 36.0, 95 per cent CI = 8.54, 151.00).

Discussion

This is the first study, of which we are aware, that examines the prevalence of all 11 DSM-III personality disorder types in a sample of cocaine abusers compared to both substance abuse and community controls. Narcissism, as assessed by the PDQ, is very infrequently found in non-cocaine substance abusers and community controls. The symptoms of narcissism reflect an emphasis on personal attention and power and are assessed in the PDQ with true responses to questions such as, "I frequently make myself the center of attention" and "I find myself thinking about how great a person I am, or will be." Cocaine abusers have often described a sense of euphoria and power with cocaine

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High Prevalence of Hepatitis B Virus Infection in Rural Black Adults in Mseleni, South Africa

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Abstract: A community-based study in northern Natal/KwaZulu, South Africa, assessed the prevalence of hepatitis B virus (HBV) infection in rural Black adults. The prevalence of HBV carriers was 4.4 percent in women (N = 342) and 7.1 percent in men (N = 99). At least one marker of HBV infection was present in 81 percent of women and 86 percent of men. The relative risk (RR) of HBV marker positivity in women due to scarification, adjusted for the presence of pierced ears, was 1.37 (95% CI = 0.9, 2.1). Risk factors such as scarification and pierced earlobes need to be further investigated to assess their role in the transmission of the HBV. (*Am J Public Health* 1989; 79:893–894.)

Introduction

Hepatitis B virus (HBV) infection has a worldwide distribution and is especially important in Africa.[1] Asymptomatic HBV carriers comprise the main infectious reservoir responsible for transmitting the virus to susceptible individuals in the population. HBV carriers also suffer higher morbidity and mortality due to chronic forms of hepatitis, cirrhosis, and primary hepatocellular carcinoma.[2]

The highest prevalence rates of HBV carriers have been reported from southeast Asia and sub-Saharan Africa. In rural Thailand, a 7.3 percent prevalence of hepatitis B surface antigen (HBsAg) has been reported[3]; in Africa the HBsAg prevalence rates vary from 2.1 percent in Kenya[4] to 12.1 percent in Somalia.[5]

In South Africa, Bersohn, *et al*,[1] observed a 2 percent prevalence of HBsAg in urban and 7 percent in rural Black South African blood donors. Prozesky, *et al*,[6] found adult HBsAg prevalences of 14.6 percent in men and 4.6 percent in women from the rural area of Kangwane. In Kavango, a rural area of northern Namibia, a 13.6 percent prevalence of HBsAg has been reported.[7] These estimates, however, are likely to be biased by self- and social-selection in samples of blood donor populations and health services users.

A previous study of 238 schoolchildren in Mseleni showed a HBsAg prevalence of 18.4 percent, one-half of whom were also hepatitis B e antigen (HBeAg) positive.[8] To our knowledge no community-based studies have been undertaken to assess the prevalence of HBV infection in rural Black adults in South Africa. The purpose of this investigation was to assess the prevalence of adult HBV carriers in Mseleni, the proportion of female HBV carriers who are HBeAg positive, and the possible role of ear piercing and scarification in the transmission of the virus.

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Methods

Field work was conducted in November 1985 in Mseleni, a rural area in the Ngwavuma district of northern Natal/KwaZulu on the east coast of South Africa. Mseleni is similar to many other rural areas in South Africa.

Each household in the district is visited regularly by field workers of the malaria control program as part of the ongoing surveillance for malaria in the area. For the purposes of this study, the typical arrangement of four to eight huts built in a circle with a common cooking area in the center is regarded as a household. Household size varies from three to 32 persons. There are approximately 1,000 such households in Mseleni, many of them being inaccessible by road. Piped water is not available. Individual households are isolated and there are often considerable distances between households.

About 60 percent of households are accessible by jeep, and these served as the primary sampling units. A multistage sampling technique was used. One adult from each accessible household was randomly selected by the field worker after he had assigned a number to each adult in the household as he encountered them during his house visit. After informed consent was obtained, the subject's name, sex, and household number were recorded and each person was examined for evidence of scarification and pierced ears. A single 5ml tube of blood was drawn for HBV studies at the conclusion of the interview.

All blood specimens were centrifuged within 24 hours, and the sera stored at −20°C and transported on dry ice to Durban (approximately 500 km away). All sera were tested by radioimmunoassay for HBsAg, hepatitis B surface antibody (HBsAb), and hepatitis B core antibody (HBcAb). Sera positive for HBsAg were further tested for IgM-class HBcAb, HBeAg and hepatitis B e antibody (HBeAb) (AUSRIA-II, AUSAB, CORAB, CORAB-M and ABBOTT-HBe; Abbott Laboratories, Chicago, Illinois). Calculations of the confidence intervals (CI) for proportions[9] and relative risk (RR) of the Mantel-Haenszel weighted risk ratio[10] and tests for statistical interaction were done by standard methods.

Results

A total of 441 subjects were studied, of whom 342 (77.6 percent) were women. The prevalence of HBsAg was 7.1 percent in men and 4.4 percent in women (RR = 1.6; 95% CI = 0.7, 3.9). At least one serological marker was positive in 81 percent of the women and 86 percent of the men (Table 1)

TABLE 1—HBV Serological Markers in Black Adults in Mseleni, South Africa

HBV Marker	Female (N = 342)	Male (N = 99)	Combined (N = 441)
HBsAg+	15 (4.4)*	7 (7.1)	22 (5.0)
HBsAb+/HBcAb+	261 (76.3)	78 (78.8)	339 (76.9)
Any Marker Present	276 (80.7)	85 (85.9)	361 (81.9)
All Markers Absent	66 (19.3)	14 (14.1)	80 (18.1)

*Number (%).

indicating previous or current HBV infection.

Only one of 15 (6.7 percent) HBsAg positive women was HBeAg positive (95% CI = 0, 28). The remaining 14 women and all seven men who were HBsAg positive were HBeAb positive and HBeAg negative. All HBsAg positive subjects were negative for IgM-class HBcAb, confirming that they were HBV carriers.

The overall response rate was estimated at between 60 percent and 70 percent. The mean age of the women studied was 43.1 years (standard deviation, 16.7 years; range, 16–88 years). Most men from the rural areas of South Africa work in the large urban centers during the year and were therefore under-represented in our study. Given this selection bias, the data on men were not further analyzed.

The prevalence of scarification was high in both HBV marker positive (96.7 percent) and negative (90.9 percent) women. The RR of being HBV marker positive was 1.36 (95% CI = 0.9, 2.1) for scarified women, taking women without scarification as the reference group. After adjustment for pierced ears, the RR of being HBV marker positive due to scarification was 1.37 (95% CI = 0.9, 2.1).

There is no evidence of positive or negative interaction between the effects of scarification and pierced ears, under a multiplicative model.*

Discussion

While a response rate of 60–70 percent is less than ideal, such a response rate is good in a rural setting which has many communication and transportation problems. Even though our sample underrepresents certain groups such as men and disabled or sick women, we believe it provides a reasonably accurate estimate of the prevalence of HBV carriers in healthy women in Mseleni.

One can infer, despite the small number of HBsAg positives, that only a small proportion of the HBsAg positive women are also HBeAg positive. The low frequency of HBeAg positivity and the high proportion of HBeAb positivity in women HBV carriers holds for women of reproductive age; the exclusion of older women does not alter the carrier rates significantly. This finding provides one explanation for the very low levels of vertical transmission in southern Africa, since the presence of HBeAg in the mother is important for transmission of HBV infection to the newborn.[11]

The high prevalence of HBeAg (50 percent) in Mseleni school girls who are HBV carriers[8] compared to the low prevalence of HBeAg in adult women (6.7 percent) suggests a high HBeAg seroconversion rate in young women. Another explanation for this difference would be a generation effect but this is unlikely considering that the HBV exposure rates in Mseleni school girls and adult women were 65 percent and 81 percent, respectively.

In southern Africa, the predominant mode of transmission is horizontal.[4] Homosexuals and intravenous drug users are relatively less important risk groups in this region.[12] The exact mechanism of transmission is unknown.

*Data available on request from author.

Both scarification[13] and pierced ears[14] have been previously implicated in the transmission of HBV. The lack of strong associations between these risk factors and HBV marker positivity in our data is possibly due to the high prevalence of scarification (95.6 percent) and pierced ears (90.6 percent). However, the upper limit of the 95% confidence interval for the RR of being HBV positive due to scarification is 2.1, which is consistent with associations reported in our previous study of Black children in South Africa.[8] Possible risk factors such as scarification and pierced earlobes need to be further investigated to assess their importance in the transmission of the HBV.

HBV infection is endemic in Mseleni. Mass vaccination campaigns to reduce the size of the susceptible population as well as health education campaigns aimed at reducing the risk of infection need to be mounted regardless of our current ignorance of specific risk factors.

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Evaluation of Radiation Safety in 29 Central Ohio Veterinary Practices

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Abstract: A sample of 29 veterinary practices in Central Ohio were visited to assess radiation safety practices and observance of state regulations. Lead aprons and gloves were usually available, but gloves were not always worn. Protective thyroid collars and lead glasses were not available in any practice, lead shields in only five practices, and lead-lined walls and doors in only two practices. Eighteen practices had none of the required safety notices posted. (*Am J Public Health* 1989; 79:895–896.)

Introduction

Exposure to ionizing radiation is a potentially serious occupational health hazard. At high doses it is carcinogenic, mutagenic, and teratogenic,[1] but the effects at low doses are much less certain.[2] The severity of the effects appears in most circumstances to be no greater than a linear relationship to dose with no threshold value. A major source of human exposure is the medical utilization of diagnostic and therapeutic radiography.[1] Veterinarians represent a subgroup of the population frequently exposed to occupational ionizing radiation. Widespread radiation exposure in the human medical profession has been minimized by specialization, but most veterinary physicians remain generalists.

The objective of our study was to assess the use of radiation safety features and the observance of state regulations in a sample of veterinary practices.

Materials and Methods

The sampling frame consisted of the 87 veterinary practices in Central Ohio having radiograph machines registered with the Radiologic Health Unit of the Ohio Department of Health. Central Ohio was defined as Franklin, Delaware, Licking, Fairfield, Pickaway, Madison, and Union Counties.

The sample size was calculated at 29 to estimate the proportion of the practices in which employees used protective equipment,[3] with an accuracy of ± 15 percent and a 95 percent confidence interval. The sample size was increased by six to compensate for potential refusals. This sample of 35 practices was chosen using computer-generated random numbers.

A letter was sent to the owners of all 35 sampled practices during July 1987. Consent was sought in a telephone interview the following week. The owners of 29 practices agreed to participate. Three owners refused to allow a site visit, another one was ineligible because the practice had been sold, and the final two were unreachable after numerous attempts over a two-week period.

The site visits were conducted during the last three weeks of August 1987. The first part of each visit consisted of a face-to-face interview about practice characteristics, personnel, and procedures used for taking radiographs. Next, a site evaluation was conducted looking at equipment, personal protection devices, and the posting of safety notices. A copy of the site visit results was left with each practice at the conclusion of the visit. The results of the site visits and interviews were analyzed descriptively.

Results

The radiation safety procedures used in the practices are summarized in Table 1. Although lead aprons and gloves were available in most of the practices, gloves were not always worn. The available aprons and gloves had been tested for leaks in only seven practices. Additionally, radiation safety training had been conducted in only 10 practices. Among the 16 practices in which film badges were worn, badges were shared by several individuals in three practices.

Collimators to narrow the x-ray beam and reduce the scatter radiation were used in 25 practices; diaphragms were used in three practices and cones were still in use at the one remaining practice. Diaphragms and cones are older methods of collimation that are no longer considered adequate to reduce scatter radiation. Techniques charts (to limit overexposure as a result of poor quality radiographs) listing the peak kilovoltage, the milliamperage, and the time of exposure necessary for taking all radiographs were available in 26 practices. Intensifying screens, which reduce the amount of exposure necessary to produce a radiograph, were available in all 29 practices. The walls and doors were lead-lined in the room containing the radiograph machine in just two practices, and in only five practices was a lead shield available for the employees to stand behind.

Three types of safety notices (the safe operating procedure, the machine registration, and a "Notice to Employees") are required to be posted near the radiograph machine according to the Ohio Administrative Code. The required safe operating procedure, which includes guidelines for taking radiographs and procedures for resolving questions and problems, was written for only eight practices. In only three practices was the safe operating procedure posted. The machine registration was posted in eight practices and the "Notice to Employees" was posted in five practices. All

TABLE 1—Summary of Radiation Safety Procedures Used in 29 Veterinary Practices in Central Ohio

Safety Procedure	% Used
Lead aprons and gloves worn	96.6
Film badges worn	55.2
Radiation safety training	34.5
Safe operating procedure written	27.6
Apron and gloves tested for leaks	24.1
Protective thyroid collars worn	0.0
Lead glasses worn	0.0

From the Departments of Preventive Medicine (Moritz, Wilkins) and Veterinary Preventive Medicine (Hueston), The Ohio State University. Address reprint requests to John R. Wilkins, III, DrPH, Associate Professor, Department of Preventive Medicine, Starling Loving Hall, 410 West 10th Avenue, Columbus, OH 43210. Ms. Moritz is currently an epidemiologist with the Commonwealth of Kentucky, Frankfort; Dr. Hueston is currently Chief Staff Veterinarian, USDA, Fort Collins, Colorado. This paper, submitted to the Journal September 13, 1988, was revised and accepted for publication February 2, 1989.

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three of the required documents were posted in only three practices; none were posted in 18 practices.

Discussion

Veterinarians come in contact with many potentially harmful substances through their work. In this study and in a companion study,* exposure to ionizing radiation did not appear to be one of the most pressing concerns for many of the veterinarians. As suggested in similar studies completed in the 1950s,[4,5] the small number of radiographs taken may account for the minimal exposure found in the companion study* rather than good protective techniques. In our study the amount of radiation protection available in the sampled practices appears to have been minimal.

Lead shields, protective thyroid collars, and lead glasses are examples of established protective equipment not frequently observed in this study. Additionally, lead-lined walls and doors in the room containing the radiograph machine would help reduce the exposure of employees working elsewhere in the facility.

Veterinarians need to strive for exposures as low as reasonably achievable (ALARA), a concept first proposed by the National Committee on Radiation Protection in 1954. This is an extremely important concept since it is generally believed that there is no threshold value for exposure to ionizing radiation. To attain this goal, protective equipment needs to be made available, maintained in good condition, and used properly and routinely. Additionally, the importance of adequate radiation protection should be continuously re-emphasized by veterinary professional organizations and colleges of veterinary medicine. Finally, regulatory agencies in each state need to ensure that all established guidelines for minimizing occupational exposure to ionizing radiation are publicized and enforced.

*Moritz SA, Hueston WD, Wilkins JR III: Patterns of Ionizing Radiation Exposure among Female Veterinarians. (Manuscript in preparation.)

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New Rule Governs Graduate Fellowships for Nursing School Faculty

A rule governing grants for postbaccalaureate fellowships for nursing school faculty was proposed in the June 5 *Federal Register* by the Health Resources and Services Administration of the US Public Health Service.

Approximately 1,440 public or private nonprofit schools of nursing would be eligible to apply for grants to enable nursing faculty to:

- Investigate cost-effective alternatives to traditional health care modalities with special attention to the needs of at-risk populations;
- Examine nursing interventions that result in positive health status outcomes, particularly with regard to family violence, drug and alcohol abuse, women's health, adolescent care, and disease prevention; or
- Address other nursing practice areas considered by the HHS secretary to require additional study.

To be eligible for fellowship support under the proposal, a faculty member must expect to meet requirements for a master's or doctoral degree before the end of the grant's budget period and be conducting a substantial study, thesis, or dissertation in a specified area. The rule proposes that grants would be awarded first to support applications demonstrating special factors related to national needs, as announced in the *Federal Register*. Approximately $1.1 million was appropriated for the postbaccalaureate faculty fellowship program in fiscal 1989.

For further information, contact Blake Crawford of the US Public Health Service. Tel: (301)443-3376.

New from nchs

Mortality among Children and Youth

Lois A. Fingerhut, MA, and Joel C. Kleinman, PhD

Introduction

Substantial decreases in the mortality of children and youth in the United States have been observed since 1900. While decreases in mortality occurred from birth to old age, children have experienced the largest declines. However, a recent National Center for Health Statistics (NCHS) report shows that despite the successes that have been achieved, there still remain appreciable age, race, and sex differences in the amount of the decline and in the current (1983–85) status of the mortality of children and youth.[1]

The large proportion of deaths from external causes—injuries, homicide and suicide (ICD codes E800–E978)—distinguishes mortality at ages 1–19 from mortality at other ages. In the early 1930s, less than one-fourth of all deaths in this age group were associated with external causes; by 1983–85, one-half of all deaths (ages 1–19) were due to injuries with an additional 14 percent attributed to homicide and suicide. Three in 10 deaths were caused by motor vehicle injuries, a proportion far greater than that for any other single cause of death in this age group.

All of the US mortality data in this report are for the 50 States and the District of Columbia and are based on data from death certificates provided by State and city vital statistics offices to the National Center for Health Statistics (NCHS) (or to its predecessor agencies). Data for eight other industrialized nations are also included, based on tabulations provided by the World Health Organization, individual country reports, and selected detailed studies.

Trends

The first half of the 20th century witnessed very large declines in mortality, especially among the youngest children, ages 1–4 and 5–9. From 1900–50, the death rate declined by about 90 percent in each of these age groups to 139 per 100,000 children ages 1–4 and to 62 per 100,000 children ages 5–9. From 1950–85, these large declines continued, reaching 51 per 100,000 among children ages 1–4 and 25 per 100,000 in children ages 5–9. Declines in mortality for children ages 10–14 and 15–19 were around 80 percent during the first half of the century. For children ages 10–14 the decline also continued—falling 50 percent during 1950–85 to 28 per 100,000 children. Among teenagers 15–19, the total decline since 1950 has been only about 25 percent; the death rate in 1985 (81 per 100,000) was only about 10 percent lower than it was during the early 1960s (Figure 1).

In 1900, external causes of death were responsible for about 10 percent of the deaths of children and youth. By 1950, however, the proportions were much higher, ranging from 27 percent at ages 1–4 to 48 percent at ages 10–19. In 1985, external causes comprised 64 percent of all such deaths with proportions ranging from 45 percent among preschoolers ages 1–4 to 78 percent among teenagers 15–19.

Current Status, 1983–85

In 1983–85, the death rate of Black children ages 1–4 was 70 percent higher than that of White children; for Black children ages 5–9 the rate was 50 percent higher. External causes of death accounted for half of the race difference in death rates for children 1–4 years of age and three-fourths of the difference for those ages 5–9 years. In particular, three causes of death (fires, homicide, and pedestrian-related motor vehicle injuries) accounted for about 45 percent of the excess deaths among Black children ages 1–4 and 65 percent of the excess among Black children ages 5–9 (Figure 2).

At ages 10–14 years, the Black-White mortality ratio decreased to 1.3 and among teenagers 15–19 the death rates for Whites and Blacks were essentially the same, a feature unique to this age group. However, there were extremely large racial differences in the causes of death for teenagers

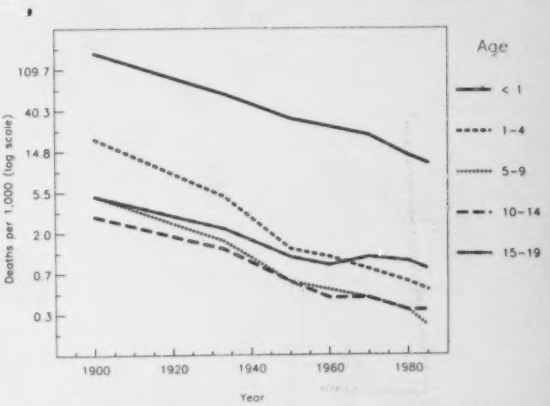

FIGURE 1—Mortality Trends among Children and Youth 1900–85

Address reprint requests to Lois A. Fingerhut, Division of Analysis, Office of Analysis and Epidemiology, National Center for Health Statistics, 3700 East-West Highway, Hyattsville, MD 20782.

Copies of the full report *Trends and Current Status in Childhood Mortality, United States, 1900–85* by L.A. Fingerhut and J.C. Kleinman, *Vital and Health Statistics* Series 3, No. 26 are available from the US Government Printing Office, Washington, DC 20402. Stock no. 017-022-01063-1, price is $2.75.

For more information about this and other reports and data from the National Center for Health Statistics, write to the Scientific and Technical Information Branch, NCHS, 3700 East-West Highway, Hyattsville, MD 20782, or call (301) 436-8500.

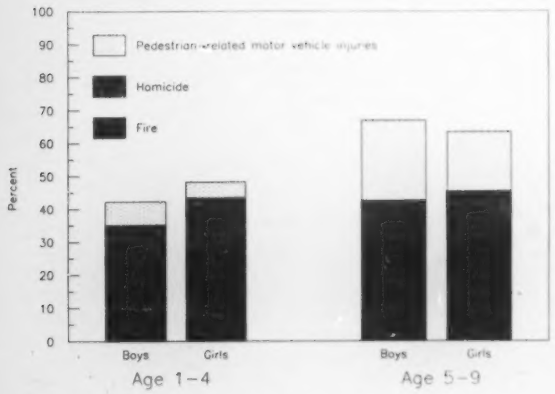

FIGURE 2—Percent of Excess Black Deaths Attributable to Specific Causes, 1983–85

15–19. The homicide rate for Blacks was five times that for Whites while the motor vehicle injury death rate for Whites was nearly three times that for Blacks.

Cross-National Comparisons, 1985

Childhood mortality in the US is generally higher than mortality in eight other industrialized nations.[2–5] However, death rates for natural causes showed little cross-national variation; natural-cause death rates in the United States generally ranked mid-way among the rates in the eight comparison countries (Table 1). Malignant neoplasms and congenital anomalies were the leading natural causes of death among children in the United States and in all of the other countries.

In contrast, US death rates for external causes ranked highest or second highest among these countries (Table 1). The causes of fatal injuries in European countries and other developed nations were similar to those in the United States—motor vehicle injuries and drownings. However, injury death rates were substantially higher in the United States, Canada, and Australia than in most of the other comparison countries. Generally, the US death rates were 1.5–2.0 times the rates in the five European countries.

Suicide rates among 15–19-year-old males varied widely in 1985, from a low of 4.3 per 100,000 in England and Wales to 18.4 in Canada. The US rate was 16.0. Among females, the rates were much lower, ranging from 1.1 in England and Wales to 5.3 in Sweden. The US rate was 3.7.

Homicide rates among 15–19-year-olds reveal a very different picture. In the United States, the rate was 13.2 per 100,000 for males, nearly five times higher than the rate in Canada and Australia and about 20 times higher than in the European countries. The excess homicide mortality in the United States is evident for both White and Black males. The homicide rate for US White males ages 15–19 (7.3 deaths per 100,000 population 15–19) was 3–24 times higher than the rates in the other countries and the rate for US Black males (46.4) was 17–155 times higher. The cross-national differences were smaller among females, but the US rates were by far the highest.[3–5]

In the United States, 74 percent of the 1,223 deaths from homicide among 15–19-year-old males (69 percent among Whites and 80 percent among Blacks) were caused by firearms (ICD code E965). Among the comparison countries, only in France did the proportion approach this level, at 12 out of 18 deaths. Forty-two percent of the homicide deaths in Canada and 31 percent of those in the Federal Republic of Germany were caused by firearms. None of the homicide deaths in Japan, England and Wales, or in Sweden were caused by firearms (Table 2). For females in the United States, 45 percent of the 356 homicide deaths were firearm-related (44 percent among Whites and 48 percent among Blacks). Similar to males, the number of these deaths in the other countries was extremely small. (Data were not readily available for Australia or The Netherlands.)

Final 1986 Mortality Data

When the Vital and Health Statistics report[1] went to

TABLE 1—Death Rates for Natural and External Causes among Children and Youth 1–19 Years of Age, by Country, Age, and Sex, United States and Selected Countries, 1985

Country	1–4 years		5–9 years		10–14 years		15–19 years	
	Male	Female	Male	Female	Male	Female	Male	Female
Natural Cause Deaths per 100,000 Population								
United States	30.9	25.9	12.2	11.4	12.1	10.6	20.4	14.9
Federal Republic of Germany	30.6	29.3	11.3	12.7	10.7	10.4	19.8	13.0
France	31.0	28.7	15.1	12.4	13.2	11.5	26.4	16.9
The Netherlands	31.8	29.6	11.7	10.4	13.7	8.7	26.1	13.9
England and Wales	35.9	32.1	12.5	13.0	13.9	12.9	24.1	15.3
Sweden	23.3	24.6	12.4	5.5	10.2	9.9	14.9	10.0
Canada	26.8	23.9	11.1	9.5	12.0	10.7	23.8	15.2
Japan	32.3	27.6	12.8	10.1	12.0	10.2	19.0	12.6
Australia	30.9	25.5	13.9	11.3	14.0	10.0	24.1	15.0
External Cause Deaths per 100,000 Population								
United States	27.2	18.6	16.0	9.8	22.8	9.9	94.3	31.6
Federal Republic of Germany	16.1	12.9	13.1	9.4	12.0	6.2	63.6	20.9
France	18.4	12.6	13.2	8.7	16.4	7.0	62.9	22.4
The Netherlands	13.7	7.9	9.4	3.2	8.5	5.2	34.1	11.0
England and Wales	13.9	8.7	9.1	5.1	15.0	5.8	43.5	12.9
Sweden	6.7	3.8	7.2	5.0	7.0	5.2	46.7	20.8
Canada	21.0	13.4	15.3	9.2	19.2	9.6	77.9	26.9
Japan	23.4	14.0	13.8	5.2	7.9	2.9	50.8	11.1
Australia	25.7	19.5	15.7	8.6	14.7	6.0	87.4	26.7

SOURCE: Data for countries other than the US are from special tabulations provided by the World Health Organization.

TABLE 2—Firearm-related Homicide Deaths among Males and Females Ages 15–19, United States and Selected Countries, 1985

Country	Males		Females	
	Number of Homicide deaths (ICD E960–E969)	Caused by Firearms (ICD E965)	Number of Homicide deaths (ICD E960–E969)	Caused by Firearms (ICD E965)
United States*	1223	903	356	161
White	556	382	201	88
Black	634	506	141	68
Federal Republic of Germany	16	5	33	5
France	18	12	3	2
England and Wales	11	0	6	0
Sweden	2	0	1	1
Canada	26	11	15	5
Japan	14	0	14	1

*In the US, homicide generally includes "legal intervention" (ICD codes E970–E978). In 1985, there were 22 deaths among males and one death among females, thus bringing the total number of homicide deaths to 1,602.
SOURCE: Data are from the 1985 annual vital statistics reports of each listed country.

press, the 1986 mortality data were not yet available. With their current availability,[6] it is worth noting that, in general, childhood and youth mortality did not change much between 1985 and 1986. However, there are specific changes that merit careful attention:

Homicide Mortality

While homicide has been the leading cause of death among Black males ages 15–19, the death rate had been declining during the years 1980 through 1984. In 1985, the rate increased from 39.3 to 46.4 per 100,000; and in 1986 the increase continued to 51.5. Thus, there was a 31 percent increase in two years.

Motor Vehicle Injury

Motor vehicle injury is the leading cause of death among White males ages 15–19. From 1981 to 1985, the death rate had declined an average of 5 percent per year. In 1986, however, the death rate increased 13 percent to 58.4, higher than it had been since 1981. Motor vehicle injury mortality increased for each of the three other race-sex groups as well. Among Black males, for whom this is the second leading cause of death, the rate increased slightly from 21.2 in 1984 to 21.9 in 1985; and in 1986 the rate increased 17 percent to 25.6, to the level it had been in the late 1970s.

Injury prevention in childhood has been targeted by the US Public Health Service as a priority concern.[7] Given the far lower death rates observed elsewhere in the world, the goals for reducing mortality from injuries, homicide, and suicide would seem to be achievable.

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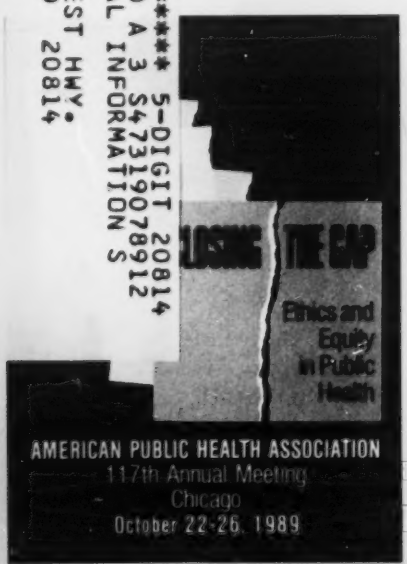

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Model Standards Impact on Local Health Department Performance in California

CAROL SPAIN, MPH, EILEEN EASTMAN, MA, AND KENNETH W. KIZER, MD, MPH

Abstract: An evaluation of a state-local negotiation process to implement Model Standards in 18 California local health departments was conducted during 1985–87. Model Standards specific to California were developed by state and local public health officials for seven programs. Nine counties participated in state-local negotiations to set performance standards in these seven programs.

During the two-year effort, greater use of Model Standards and more improvement in local program performance occurred in local health departments that participated in the state-local negotiation process. The use of Model Standards appeared to contribute to establishing program priorities, emphasizing the measurement of outcomes, improving data management systems, and evaluating the current performance of programs. While this state-local negotiation process is suggested as the preferred approach to encouraging the use of Model Standards in local health departments, it must be recognized that this process requires time and commitment of resources from both the state and local levels. (*Am J Public Health* 1989; 79:969–974.)

Introduction

Experience utilizing the Model Standards has substantially increased since 1979, when their use began as a collaborative effort among the US Conference of City Health Officers, the National Association of County Health Officials, the Association of State and Territorial Health Officials, the American Public Health Association (APHA), and the federal Centers for Disease Control (CDC).[1] Beginning with the publication of *Model Standards for Community Preventive Health Services*, a number of reports have described uses for or experience with public health standards in general and with Model Standards in particular.[2–9]

Three major approaches have been used to implement the Model Standards process, each requiring different levels of administrative sponsorship: 1) statewide projects that have had strong state health department sponsorship; 2) local health department actions without state sponsorship; and 3) individual efforts by local program managers without strong local or state health department sponsorship. The relative effectiveness of the different levels of sponsorship has not been determined. However, the APHA and CDC are currently conducting a Model Standards Technical Assistance Project which is supporting and documenting the progress made in 13 local and state Model Standards projects.[10]

Two evaluations of the use of Model Standards have been conducted in California.[11,12] The California approach to implementing Model Standards has involved the use of a state-local negotiation process to set local Model Standards objectives. On reporting the first project results, Weiler stated, "Project findings suggest that the Model Standards negotiations could provide state and local levels of government with a valuable tool for determining health care priorities and generating objective programmatic data for budget justification." The strength of the project was determined to be the negotiation process itself, but both state and local program managers felt that the initial limited evaluation period did not allow for adequate assessment of the results of the process.

This report describes the results of more extensive efforts to assess the impact of Model Standards implementation in California. Specifically, this evaluation looked at five areas: 1) setting of objectives via a state-local negotiation process; 2) achievement of objectives and improvement in program performance; 3) impact of data availability and improvements in data collection; 4) administrative evaluation of the Model Standards process; and 5) key determinants to successful application of Model Standards.

Methods

There were two major components of this effort: 1) the specification of Model Standards specific to California, and 2) a comparison of the impact of state-local negotiations over the two-year observation period, January 1985 through December 1986. California-specific Model Standards were established for seven programs: Chronic Disease Control, Sexually Transmitted Disease (STD) Control, Tuberculosis Control, Food Protection, Nutritional Services, Vehicular Injury Control, and Non-Vehicular Injury Control. Eighteen counties participated, representing the demographic and geographic diversity of the State. The Health Officers Association of California (HOAC), a private nonprofit membership organization of California's local health officials, was the independent organizational subcontractor charged with observing and evaluating the project.

Standards Specification

The first component of this project involved the specification of Model Standards for the seven programs noted above. A goal statement, outcome and process objectives, and key data indicators were specified for each of the seven programs. A total of 94 objectives were specified. The document used as the basis for this specification was the first edition of *Model Standards*.[1] Committees comprised of both state and local program managers and health officers met over a three-month period to develop these standards, which were approved by the California Conference of Local Health Officers in April 1984.

Selection of Counties

California has 61 local health departments (56 single county health departments, one bi-county health department,

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and four city health departments). State law delegates significant authority to the local level, while retaining broad state enforcement powers and considerable public health activity at the state level. Local health officers have a strong working relationship with the California Department of Health Services and are statutorily represented by the California Conference of Local Health Officers (CCLHO). HOAC conducts public health legislative, research, and educational activities.

Eighteen county health departments participated in this project; nine of them participated in negotiations with the State. All were selected because of their willingness to participate as *either* a "negotiating" or "nonnegotiating" county and their fitness to be paired with a comparable county for evaluative purposes. To select pairs representative of the State, a number of variables were identified as important selection criteria, and counties were paired according to these attributes (See Appendix).

Setting of Objectives

Prior to the beginning of each of the two observation years, all 18 counties completed an assessment document that included the California-specific Model Standards for the seven programs in the project. Each program standard contained a goal statement followed by a series of program objectives. For each program objective the following questions were asked:

- What is the current status of the objective?
- When can the objective be achieved?
- To what degree can the objective be achieved within the next year?
- How can achievement of the objective be measured or demonstrated?

In addition, State Health Department staff in the seven programs used the same assessment documents to assess the project counties from the State's perspective.

Negotiations were then undertaken with nine counties. Completion of the negotiation process for all seven programs required two days, with each program negotiation session lasting approximately two hours. Participants in each negotiation session generally included the local health officer, local program manager, state program representative, a facilitator, and an observer. Usually, the state and local program representatives were the same individuals responsible for completing the assessment documents. Negotiations were on-site in the local jurisdiction so that local staff participation could be maximized. In all, 125 separate negotiation sessions were conducted during the two years, one at the beginning of each year in each of the nine counties.

In the nine counties not participating in the state-local negotiation process, program objectives were set by the local health department.

Local Implementation and Evaluation

Progress toward achieving the objectives set was evaluated at the end of each of the two years by means of on-site interviews and mailed questionnaires. Data were gathered to document changes in program performance, staff, funding, and data management.

Improvement in individual program performance was one of the key outcomes evaluated. Improvements in program performance were equated with setting and achieving program objectives (both process and outcome) at higher levels of performance than the baseline performance of the program as determined by the state and local health departments. Improvements in program performance were also self-reported in general by program managers and, if related to the use of Model Standards, were noted as such. No external measurement of program performance or improvement in program outcomes was conducted.

Results

Setting of Objectives and State/Local Negotiations

Ninety-four objectives covering all seven program areas were specified at the onset of the project so that, in the aggregate, a total of 1,692 objectives (846 each year) could have been set for both the negotiating and nonnegotiating health departments. During the two-year assessment, negotiating health departments set a total of 1,160 objectives, and nonnegotiating health departments set 830 objectives. A drop in target objectives occurred in the second year, with nonnegotiating health departments decreasing the number of objectives by 70 percent (637 to 193) and negotiating health departments by only 10 percent (611 to 549).

Negotiating health departments specified an increased level of performance in 725 objectives over two years, while this occurred for only 223 objectives in nonnegotiating health departments (Table 1).

In both groups, medium- and small-sized counties set their objectives at higher levels of performance most frequently. The specific programs in the second year that set objectives at higher performance levels most frequently were Non-Vehicular Injury Control (62 percent), Nutrition Services (62 percent), and Chronic Disease Control (60 percent). These programs generally had not achieved high levels of program performance prior to the study and had many objectives yet to achieve.

In both years, negotiating health department staff rated the state-local negotiation process as the most useful aspect of the project. Eighty-two percent of the programs reported staff benefits from negotiations, and 79 percent of the programs reported that state-local negotiations had a positive impact on program performance.

Achievement of Objectives and Improvements in Program Performance

As shown in Table 2, both negotiating and nonnegotiating departments achieved similar proportions of the objectives they had set but negotiating departments achieved 507 objectives and their counterparts only 177. Control of Sexually Transmitted Diseases and Vehicular Injury were the programs for which both sets of health departments reported the most objectives achieved.

The programs whose improvements in program performance were considered by both groups as due to the use of Model Standards were Tuberculosis Control (61 percent), STD Control (61 percent), and Nutrition Services (50 percent). Negotiating health departments reported that these programs also received significant technical assistance from state staff as a by-product of the negotiation process.

Negotiating health departments reported that Model Standards assisted in improving program performance in 40 of 63 programs, in contrast to 15 to 63 programs in nonnegotiating health departments. Programs that had a program organizational unit (e.g., Tuberculosis Control) and those that had state funding reported this impact most frequently. Thirty-two programs in both groups reported staff increases during the two-year study and 11 of these reported that the increases were due in part to the use of Model Standards. These staff increases were achieved primarily through inter-

TABLE 1—Setting of Objectives[a]

	Negotiating Health Departments		Nonnegotiating Health Departments
	Prior to Negotiation	Post Negotiation	
Year 1			
Health department agrees to increase its level of performance	19% (N = 97)	72% (N = 440)	19% (N = 121)
Health department agrees to do the same as or less than the previous year[b]	81% (N = 411)	28% (N = 171)	81% (N = 516)
Totals	N = 508	N = 611	N = 637
Year 2			
Health department agrees to increase its level of performance	36% (N = 141)	53% (N = 285)	52% (N = 102)
Health department agrees to do the same as or less than the previous year[b]	64% (N = 250)	48% (N = 264)	47% (N = 91)
Totals	N = 391	N = 549	N = 193

[a]There was a total of 94 objectives specified for the seven program areas. Thus the total possible number of objectives was 846 per year for each set of health departments (negotiating and nonnegotiating).
[b]Objectives that were set at "maintenance of effort or less" levels were considered "pre-achieved."

TABLE 2—Achievement of Objectives[a]

	Number of Objectives		
	Objectives Achieved[a]	Objectives Not Achieved	TOTAL
Year 1			
Negotiating	65% (N = 282)	35% (N = 152)	100% (N = 434)
Nonnegotiating	59% (N = 71)	41% (N = 49)	100% (N = 120)
Totals	63% (N = 353)	37% (N = 207)	100% (N = 560)
Year 2			
Negotiating	64% (N = 225)	36% (N = 182)	100% (N = 507)
Nonnegotiating	63% (N = 106)	37% (N = 63)	100% (N = 169)
Totals	63% (N = 431)	36% (N = 245)	100% (N = 676)

[a]This refers to objectives set at higher performance levels and objectives set at the same level (maintenance) but achieved at a higher level.

nal redirections in response to the commitment to achieve a higher level of performance.

Impact of Data Availability and Improvements in Data Collection

Overall, 15 of the 18 health departments reported that data availability was an important consideration in their decision to set specific Model Standards objectives. TB and STD Control programs had the most readily available data. In contrast, the setting of objectives for Non-Vehicular Injury Control and Chronic Disease Control was hindered by the paucity of data to measure outcomes.

Eight of the nine negotiating health departments reported that Model Standards had a positive impact on data management, in contrast to four of the nonnegotiating health departments. Most data improvements were in large county health departments, and ranged from gathering new primary data (e.g., smoking prevalence among pregnant women) to automating existing data systems. Data-related factors that contributed to the successful use of Model Standards included:

- knowledge about data sources that already existed;
- commitment to allocate staff time to analyze existing data;
- Ideas generated by local staff about additional existing data sources and types of new primary data that could be easily collected;
- capacity to automate the local health department data system; and
- knowledge and experience of local program managers in using outcome data in program planning and evaluation.

Administrative Evaluation of the Model Standards Process

Model Standards were used primarily by negotiating counties as an administrative tool to increase program accountability and responsibility for achieving objectives; to improve evaluation of current program effectiveness and efficiency; to measure results (outcomes); to enhance communication between program managers and their staff; and to provide an opportunity for staff to participate in decisions that directly affect their work. Other uses included improving communication between the local and state health department and generally improving the state-local relationship. Major barriers to the use of Model Standards reported from both sets of counties were lack of staff, lack of program priority status within the health department, and lack of data.

Ongoing Use of Model Standards

Two of the major indicators of the usefulness of Model Standards are the degree to which the negotiating health departments have integrated Model Standards into ongoing planning and evaluation documents and activities and their intention to continue to use Model Standards. In both areas negotiating health departments responded more positively than nonnegotiating health departments. For example, only one of nine nonnegotiating counties reported integration of Model Standards, while five of nine negotiating health departments reported such integration. Only eight of 63 programs in nonnegotiating health departments intended to use Model Standards in the future, in contrast to 30 programs in negotiating counties.

Key Determinants to the Successful Application of Model Standards

Several local factors determined the success of using Model Standards in local health departments. These included:

- *Strong Local Health Officer and Program Manager Commitment.* Commitment to use Model Standards by the health officer was rated as extremely important to the successful use of Model Standards by most participants. In counties where the program manager was very involved, both in completing the assessment and in participating in the state-local negotiations, Model Standards were observed to be more useful and there were more improvements noted in program performance and data management.
- *The Priority Status of the Program within the Local Health Department.* The seven programs in which Model Standards were used were pre-selected by the State Health Department based on criteria set by CDC in their proposal guidelines. The same set of programs was used in all 18 counties. In order to be most successful, however, it appears as though local health departments should select their own priority programs to which they apply the Model Standards. For example, many local health departments felt that Chronic Disease Control and Injury Control were not their highest priorities, and they set fewer objectives in these areas.
- *Availability of Data to Set and Measure Objectives.* Health departments that were unable to gather baseline and program activity data had the most difficulty with setting and measuring objectives because they had neither reference points nor a way to measure achievement. This seems to be an area that must be effectively addressed at both state and local levels if Model Standards are to be used effectively.
- *Relationship of Model Standards to Existing Planning and Evaluation Tools.* Many local health departments were setting program process objectives and linking their program plans to state and local funding before the publication of the first edition of *Model Standards*. From this assessment, it appears as though the most successful health departments have used the Model Standards as a more specific and technical guide for program objective setting and integrated their "new" program objectives (based on Model Standards) into required county or state mandated plans tied to local budgets and contract funding. In health departments where Model Standards were linked to existing planning requirements, program managers reported improvements in the presentation of their program plans and budgets and, in some cases, increased funding for programs.

Discussion

Overall, it appears that use of a state-local negotiation process results in greater use of Model Standards and more improvement in program performance by local health departments than in the absence of such a negotiation process. Negotiations provided a forum for state and local health department staff to come together, share data and program information, and set realistic and meaningful objectives. The negotiation process improved over time because of an observed "learning curve." Both state and local negotiators understood the scope of the objectives and the indicators that existed, or could be developed, to measure outcomes, and they were more skilled at representing their positions.

We believe it important that the first step in this negotiation process be a consensus conference or series of conferences at which state and local health officials reach agreement on state-specific Model Standards. The time and effort necessary to achieve this consensus should not be underestimated, and our experience indicates that this requires clear and vigorous sponsorship at both the state and local levels.

In our project, technical assistance given to the local health departments by the state negotiators was not tied to funding, contract negotiations, or contract monitoring and it was rated very highly by local participants—possibly due to this factor. The discussions between state and local participants during the negotiations seemed to help the local participants adapt the standards to their resources, programs, priorities, and data availability. Much of the technical assistance was simply a by-product of the negotiation process itself. Although negotiating health departments had the same concerns about their limitations in setting and achieving the objectives as nonnegotiating health departments, they often chose to set more objectives because the negotiation discussion helped to focus on what a local health department could do within a one-year period to achieve a particular objective.

Both state negotiators and negotiating local health departments rated the negotiation process highly. The state negotiators generally felt that the ability to offer broad programmatic assistance was stimulating to themselves and useful to local staff. They also achieved a greater understanding of local programs, a better idea of the types of state assistance needed by counties to set local standards, (e.g., data) and an appreciation for which local health departments needed technical assistance in the future.

There are a number of difficulties that are encountered when conducting an evaluative study such as this one, including difficulties in achieving a true "control." In reality, the interest of local health departments in participating in the study was a factor that had to be considered. Although more counties were interested than were paired and participated, those that participated, either in the negotiating or nonnegotiating groups, were interested in being exposed to the Model Standards process. The true controls, one could argue, were the county health departments that did not exhibit an interest in participating in the study. In addition, because there was no external measurement of program performance conducted in either group of counties, the achievement of objectives and assessment of improvements in program performance, staffing, and funding were self-reported. Some of the achievements were related to improved outcomes in individual programs (e.g., more individuals served; delivery of a new program component) and some were related to process achievements (e.g., number of objectives set; integration of Model Standards into planning efforts).

From our experience, we conclude that a state-local negotiation process does improve local health department program performance but requires time and commitment of resources from both state and local health departments. We also believe such a negotiation process is the preferred approach to encouraging the use of Model Standards by local health departments. However, to be meaningful and to optimally meet the needs of both local and state health departments, both the use of Model Standards and the negotiation process should be integrated into the state health department's local assistance activities and should relate to clear state health department goals and administrative support.

APPENDIX
Distribution of Key Variables for Paired Counties

	SMALL COUNTIES[2]						MEDIUM COUNTIES[3]						LARGE COUNTIES[4]					
	(N)*	(NN)**	(N)*	(NN)**	(N)*	(NN)**	(N)*	(NN)**	(N)*	(NN)**	(N)*	(NN)**	(N)*	(NN)**	(N)*	(NN)**	(N)*	(NN)**
Variables[1]	Mendocino	El Dorado	Madera	Yuba	Nevada	Tuolumne	Monterey	San Luis Obispo	Santa Cruz	Yolo	Butte	Humboldt	Orange	Santa Clara	San Bernardino	Riverside	Santa Barbara	Ventura
Total County Population (1000)	67	88	64	50	53	34	291	157	188	114	144	108	1900	1300	900	700	300	500
Percent Urban	31.6	42.5	47.7	71.4	13.0	9.6	77.4	75.9	81.4	81.9	70.9	56.4	99.7	97.7	90.1	82.4	90.8	94.6
Percent Young (<5 years)	7.8	6.6	9.1	9.1	6.6	6.5	8.4	5.9	6.6	6.7	6.2	7.0	6.7	6.9	8.5	7.6	6.2	8.1
Percent 65 Years +	11.9	9.9	10.8	9.3	15.3	13.8	9.2	13.4	13.2	8.7	15.5	10.2	8.3	7.5	10.0	14.9	11.3	8.3
Percent Black	0.5	0.3	3.3	4.4	0.1	1.1	6.3	1.7	0.7	1.7	1.2	0.5	1.3	3.3	5.2	4.5	2.5	2.0
Percent Hispanic	5.5	4.6	26.8	8.8	3.0	5.2	25.9	9.5	14.7	17.1	5.2	3.4	14.8	17.5	18.5	18.8	18.5	21.4
Percent Under Poverty Level	12.3	8.7	15.7	16.1	8.7	11.9	11.4	13.7	12.2	15.9	15.0	13.0	7.3	7.1	11.1	11.3	10.6	8.0
Percent with 12+ Years Education	76.4	81.2	60.1	63.1	82.0	77.3	71.0	76.8	77.8	73.5	71.6	76.0	80.4	79.5	71.0	68.9	79.1	75.9
Median Income per Family ($1000s)	17.7	20.2	17.3	16.2	18.8	16.9	20.0	18.2	20.7	20.5	16.3	18.5	25.9	26.7	20.0	18.7	21.6	23.6
Median House Value ($1000s)	70	84	60	47	80	68	87	78	94	67	59	57	108	109	63	68	104	93
Ration of Physicians to Population (100s)	5.9	8.9	16.0	9.8	6.8	7.3	8.3	6.6	6.8	5.6	7.5	6.1	6.0	4.8	7.6	8.6	4.9	7.2
Per Capita Public Health Service Expenditures	19.7	19.2	14.7	19.5	19.1	21.8	19.2	19.6	20.8	17.7	15.4	19.1	13.1	15.8	14.9	18.6	14.1	10.3
Total Public Health Services Expenditures ($1,000,000)[5]	11.3	17.9	8.9	9.8	11.3	6.5	61.1	32.7	37.9	19.4	23.1	23.3	272.6	217.5	142.9	132.1	54.0	62.8
Number of Health Department Employees (FTE)	46.3	38.5	31.5	30.0	29.6	22.9	156.0	144.3	165.5	61.0	74.7	72.4	734.7	526.5	513.3	385.0	161.0	135.7

[1]County demographic data obtained from the 1980 census, California State Department of Finance. Figures for ratio of population to physicians obtained from the *Health Data Summaries for California Counties*, 1982. Data for per capita public health services expenditures obtained from the *County Health Services Report*, 1982. Data for total public health services expenditures and number of health department employees obtained from the *County Health Services Report*, 1983.
[2]Counties under 100,000 population.
[3]Counties 100,000–300,000 population.
[4]Counties over 300,000 population.
[5]Figures represent estimated expenditures for FY 1981–1982.
*N = Negotiating
**NN = Non-negotiating

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AIDS Triggers Increase in State Worker Notification Laws

Notifying workers of possible exposure to the acquired immunodeficiency syndrome (AIDS) virus has become an important focus of state legislative activity, according to a report published by the AIDS Policy Center of the Intergovernmental Health Policy Project. The report, based on a 50-state review of more than 100 worker notification laws enacted between 1983–88, found that, at the end of 1987, 16 states had laws on the books requiring that workers—primarily emergency medical technicians (EMTs) and funeral personnel—be notified if they were exposed to HIV (human immunodeficiency virus) or other diseases. By the end of 1988, the number had doubled. More than half of state worker notification laws mandate that both emergency care personnel and a variety of other workers, ranging from health care providers to corrections staff, be notified of actual or potential exposure to HIV.

Some laws define patient contact very broadly referring to "any contact," although more states now require that a worker must have "significant exposure"—i.e., direct contact and exchange of blood, needlesticks, and scalpel cuts—before notification will occur. Since 1988, state laws in 16 states define "significant exposure," compared to only three states before 1988.

The issue of testing following HIV exposure is a feature in a number of notification laws, although procedures regarding testing are not always outlined in these statutes. Notification leads to required testing in some states, usually of the patient to whom a worker was exposed. Patient informed consent is not always required, although 14 states now require or permit patients to be tested following worker exposure to HIV. Only six states (FL, LA, ME, RI, UT, WA) mandate patient informed consent before testing.

Few of the worker notification laws specify how the testing-after-exposure information is to be managed. A few states have levied penalties for breach of confidential information, including criminal penalties (AL, GA, IN, UT); licensure sanctions (ME, VA), dismissal from employment (IN).

Worker notification laws in 15 states require workers to be educated or trained in preventive procedures. Michigan holds employers liable in settings that have not developed and implemented policy guidelines for breaches of confidentiality of HIV-related information.

For further information about the report, contact: AIDS Policy Center, Intergovernmental Health Policy Project, George Washington University, 2021 K St, NW, Suite 210, Washington, DC 20006. Tel: (202) 676-8144.

The External Costs of a Sedentary Life-style

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Abstract: Using data from the National Health Interview Survey and the RAND Health Insurance Experiment, we estimated the external costs (costs borne by others) of a sedentary life-style. External costs stem from additional payments received by sedentary individuals from collectively financed programs such as health insurance, sick-leave coverage, disability insurance, and group life insurance. Those with sedentary life-styles incur higher medical costs, but their life expectancy at age 20 is 10 months less so they collect less public and private pensions. The pension costs come late in life, as do some of the medical costs, and so the estimate of the external cost is sensitive to the discount rate used. At a 5 percent rate of discount, the lifetime subsidy from others to those with a sedentary life style is $1,900. Our estimate of the subsidy is also sensitive to the assumed effect of exercise on mortality. The subsidy is a rationale for public support of recreational facilities such as parks and swimming pools and employer support of programs to increase exercise. (*Am J Public Health* 1989; 79:975–981.)

Introduction

A growing body of literature indicates that regular exercise may prove beneficial in ameliorating the consequences or preventing the onset of a wide range of diseases and chronic conditions.[1–9] The consequent improvements in health may translate into decreased use of health services and increased job productivity, resulting in economic benefits to society at large. Conversely, those who choose to lead more sedentary lives may actually impose costs on others.

We term such costs—costs that others pay as a result of an individual's decision to lead a relatively inactive life rather than exercise—external costs. By contrast, internal costs are those borne by the individual and are presumably taken into account in decisions about how much to exercise. External costs stem from additional payments received by sedentary individuals from collectively financed programs such as health insurance, sick-leave coverage, disability insurance, and group life insurance. The premiums and payroll taxes that finance these programs are the same for sedentary and active individuals. As a result, these social welfare programs may have the unfortunate side effect of subsidizing unhealthy behavior.

The external costs of smoking and drinking can in principle be converted to internal costs by appropriate excise taxes, but it is difficult to imagine how to tax a sedentary life-style. If, however, a sedentary life-style does impose external costs, subsidies to facilities and programs that promote exercise could well save money in the long run. Such subsidies might take the form of public parks or other recreational facilities that charge no fee or have fees below cost. The greater the external costs of inactivity, the greater the possible value of exercise promotion. In order to shed further light on this issue, we have calculated the external costs of a sedentary life-style.

Our focus on external costs can be contrasted with a recent analysis of the total (internal and external) costs and benefits of exercise.[10] Under assumptions similar to ours, exercise was shown to improve health efficiently, especially for those who enjoyed exercising. That analysis should encourage people to exercise because it is good for them. In this paper, we study a different question, which is more relevant to government action—if someone exercises, is that good for other members of society?

Methods

We will briefly describe our methods and data; more details can be found in the Appendix and in a forthcoming report, available from the authors.*

Cost Analysis

We focused our cost analysis on the lifetime (as opposed to current) costs incurred by individuals who exercise to varying degrees because some of the effects of exercise persist for many years. In judging any policy with long-term effects, it is important to discount costs that are incurred in the future. Most analysts of health policies consider a 5 percent discount rate as standard, but because the proper rate of discount is controversial, we have computed results for several discount rates that span the range commonly used.[11]

In order to assess the costs associated with reduced physical activity, we made use of a hypothetical person: the "active inactive" individual. Because we were only interested in estimating the effects of exercise on external lifetime costs, we needed to hold constant the effects of other differences between those who do and do not exercise regularly; this we did by changing only the value of the exercise variable in our estimated regression equation. Thus, we use the "active inactive" individual to determine what the lifetime costs would be if the group of inactive people in our sample had instead been active throughout their lives but retained all of their other characteristics.

Our modeling approach involved tracking, from age 20 to death, hypothetical cohorts of sedentary men and women who, for the purposes of comparison, were counterfactually made physically active. To accomplish this, we developed a life table, based on exercise status, that showed the probability of surviving to each age (see Appendix).

We used the life tables to compute the difference in costs that are due to a lack of exercise. These expected net external lifetime costs, X, are given by the formula:

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*Manning WG, Keeler EB, Newhouse JP, Sloss EM, Wasserman J: The external economic costs of poor health habits. Santa Monica, CA: The RAND Corporation, forthcoming.

$$X = \sum_{t=20}^{z} \delta^{t-20} \cdot P(A|S)_t \cdot C(S)_t$$

$$- \sum_{t=20}^{z} \delta^{t-20} \cdot P(A|RE)_t \cdot C(RE)_t$$

where

δ = annual discount factor ($= 1/(1 + r)$),
$P(A|S)_t$ = probability of surviving from 20 to at least age t given a sedentary life style,
$C(S)_t$ = annual costs minus taxes and premiums at age t for those who are sedentary,
$P(A|RE)_t$ = probability of surviving from 20 to at least age t given regular exercise
$C(RE)_t$ = annual costs minus taxes and premiums at age t for those who exercise regularly.

The lifetime external costs of leading a sedentary life included in our analysis are: covered medical costs, covered work loss, group life insurance, the widow's bonus in Social Security, covered nursing home costs, pensions, and disability insurance. (Covered means paid through a collectively financed public or private policy.) Because inactive people lead shorter lives, they will pay less of the taxes and premiums that finance these programs. This differential adds to the lifetime costs borne by active people and so must be included (see Appendix).

Although the cost analyses follow hypothetical cohorts over time, the information for what happens at each age is based on recent (cross-sectional) experience for that age, and not on projections of what life will be like, e.g., in the year 2048 when those 20 years old in 1988 turn 80. Thus, we did not estimate the effects of secular and biomedical trends on the sequalae of a sedentary life-style. Instead, we used current estimates for such parameters as retirement, medical and nursing home costs, education, and life insurance arrangement. This simplifying assumption is commonly made (e.g., by the National Center for Health Statistics[12] in computing life expectancy) because the alternative is too difficult and too conjectural.

Data

The primary data source for those under age 60 is the RAND Health Insurance Experiment (HIE), which has detailed information on health habits, diseases and chronic conditions, and the utilization of medical services. Because those aged 62 or over at the time of enrollment were excluded from the HIE sample and because we wanted to estimate lifetime costs, we used data on the elderly from a 1983 supplement to the National Health Interview Survey (NHIS) that includes information on health habits, use of medical services, and work loss. We supplemented these two sources with information from the Current Population Survey (CPS), the National Medical Care Utilization and Expenditure Survey (NMCUES), the Centers for Disease Control (CDC), and other sources.

The HIE sample is a random sample of individuals observed between November 1974 and January 1982 from six sites throughout the United States (Dayton, OH; Seattle, WA; Fitchburg, MA; Franklin, MA; Charleston, SC; and Georgetown, SC).[13,14] For this paper, we used data on 3,074 individuals enrolled in fee-for-service plans who were between the ages of 20 and 59. For those age 60 and over, we used unpublished data from the National Health Interview Survey (NHIS).[15] These data are from a supplemental questionnaire on health habits that was administered to 20 percent of the respondents to the 1983 NHIS who were at least 18 years of age; this yielded a sample of 22,418 persons.

Measures of External Costs

The costs of medical services are external to the degree they are financed by health insurance. We know the fraction of inpatient and outpatient expense paid out of pocket on average for each age, so it suffices to compute the effects of exercise on total medical spending. In our analysis, we examined differences in medical expenditures between those who engaged in regular exercise and those who did not. In the case of the HIE, medical expenditure measures were based on insurance claims filed during the experiment for periods up to five years; and in the case of the NHIS, on reported outpatient visits and hospital admissions in the preceding year.

The work loss measure used in the analysis for those under age 60 is based on responses to a health diary that was filled out biweekly by HIE participants. Work missed by those surveyed in the NHIS was estimated using responses to a question concerning the amount of work missed over a two-week period. Data on wages and pensions come from the Current Population Survey.[16,17]

Measures of Exercise Status

We classified each person in the HIE data set into one of the following three categories based on his or her reported level of exercise (ascertained by a single question) at the time of enrollment into the study: those with light or no exercise (mostly sitting or walking); moderate or strenuous exercise several times a week; or strenuous exercise most days. Individuals with role or physical limitations were separated from the rest of the population to avoid attributing the effects of their health problems to lack of exercise.

The exercise categories for the NHIS sample are also based on responses to a single question: "Are you less active, about as active, or more active than others your age?" Again, we placed individuals with physical or role limitations in a separate category. Thus, the NHIS categories differ from those of the HIE in that the NHIS measures each person's perceived amount of exercise relative to the average for his or her age group, rather than the more objective HIE measure of how often the person exercises.

Other Covariates

Our estimates of the effect of exercise on the costs of medical care and sick leave controlled for age, sex, race, smoking, drinking, education, health status, family income, and family size by including them as explanatory variables in a multiple regression analysis. In the analysis of the HIE data, we also controlled for coinsurance rate and the use of seatbelts. Education and seatbelt use are proxies for differences between those who engage in regular exercise and those who do not that may affect work loss and the use of medical services but are not affected by the amount of exercise.

Results

Use of Medical Services and Work Loss

Although moderate exercisers had 12 percent fewer outpatient episodes than the group with little or no exercise, strenuous exercisers had only 8 percent fewer outpatient episodes than those with little exercise (Table 1). The effect of strenuous exercise was substantially larger for inpatient use than for outpatient care (30 percent lower inpatient use

TABLE 1—Differences in Medical Use and Work Loss in the Health Insurance Experiment (HIE) Sample

Level of Exercise	Outpatient Episodes (excluding well-care)		Inpatient Admissions (excluding maternity)		Male Work Loss[a]	
	Ratio	95% CI	Ratio	95% CI	Ratio	95% CI
Little	100	—	100	—	100	—
Moderate	88	(80, 96)	90	(71, 113)	83	(61, 112)
Strenuous	92	(81, 105)	72	(52, 100.2)	68	(47, 99)

NOTE: Entry indicates percentage ratio of use to use by those with little exercise with similar non-exercise characteristics.
[a]HIE data did not permit an estimate of work loss for females.

rate compared to a 8 percent lower rate of use of outpatient care).

We were concerned that health status could affect exercise, as well as exercise affecting health status—the syndrome expressed by, "I exercise if I feel well." Even though we controlled for role or physical limitations, we felt that perhaps our dichotomous limitation measure was too crude a measure of ability to exercise. Hence, to check the sensitivity of our results to the possibility that health status may affect exercise, we added general health and mental health indices and a count of the number of chronic diseases to our list of covariates.[16] Strenuous exercisers had a 20 percent lower hospitalization rate and used 6 percent less outpatient care than light exercisers, after these adjustments (results not shown). Thus, the magnitude of the effect of exercise on inpatient use depends on how one treats the causal relationship between health status and exercise.

For adult males, moderate exercisers had 18 percent fewer work loss days than those with little exercise, while those with strenuous exercise missed 32 percent fewer days.

We also examined the effects of lack of exercise using the 1983 NHIS for all adults (Table 2). The NHIS sample exhibited a more pronounced effect of exercise than the HIE sample. As an example, for moderate exercisers, the HIE data exhibited a 12 percent lower outpatient use rate than for never or light exercisers, in comparison to the 29 percent lower outpatient use rate in the NHIS data for moderate exercisers.

Cost Analysis

As a base case, we estimated the external costs of those who, although not physically limited, do not exercise even when young—about 12 percent of men and 20 percent of women in our population. The lifetime external costs in each category for those who classified themselves as not very

TABLE 2—Differences in Medical Use and Work Loss in the National Health Interview Survey (NHIS), 1983

Level of Exercise	Outpatients Visits		Hospital Admissions		Work Loss	
	Ratio	95% CI	Ratio	95% CI	Ratio	95% CI
Below average	100	—	100	—	100	—
Average	71	(65, 78)	66	(58, 76)	70	(51, 95)
Above average	72	(65, 80)	65	(57, 75)	77	(56, 100)

NOTE: Entry indicates percentage ratio of use to use by those with little exercise with similar non-exercise characteristics.

TABLE 3—An Inactive Person's Lifetime External Costs

	Discount Rate		
	0%	5%	10%
Costs[a]			
Medical care[b]	60	11	5
Sick leave	5	2	1
Group life	4	1	0.4
Nursing home	11	0.7	0.1
Retirement pension	129	16	4
Taxes on earnings[a]	183	27	9
Total net costs[a,c]	27	4	1
Life expectancy at age 20 (years)	56.9	21.2	13.0

[a]Measured in thousands of 1986 dollars.
[b]Excludes maternity and well care.
[c](Sum of costs) minus taxes on wages. For example, 27 = 60 + 5 + 4 + 11 + 129 − 183. Due to rounding, categories may not sum to total.

active physically are given in Table 3. For those 20–59, we used data from the HIE and included all medical expenses (except maternity and well care) and all covered work loss; for the aged, we included all medical expenses based on data from the NHIS.

For undiscounted lifetime costs, medical costs and retirement pensions are the largest external costs, and after subtracting taxes on earnings, the total net undiscounted costs are $27,000 per person (Table 3).

Costs discounted at 5 percent show much lower nursing home and pension costs. The other costs fall less because a part of them occurs early in adulthood. The discounted lifetime external costs fall to $4,000 at a 5 percent discount rate, and $1,000 at a 10 percent rate.

Table 4 shows the change in these costs if relatively inactive individuals had exercised regularly but otherwise retained their characteristics and habits (i.e., if they became "active inactive" people). As discussed in the Appendix, we modify the CDC Health Risk Appraisal (HRA) estimates of the effects of exercise on men, and combine recent evidence[5,6,10] in assuming exercise adds 1.5 years of life for men. Then, additional exercise increases total life expectancy by 300 (undiscounted) days, and reduces medical costs. Because active individuals live longer, nursing home payments increase as do pensions. Overall, however, the effect of increased exercise is to decrease undiscounted external

TABLE 4—Difference between External Costs of Relatively Inactive Persons and Those Same Individuals Had They Been Active

	Discount Rate		
	0%	5%	10%
Differences in Costs[a]			
Medical care[b]	14.8	1.8	0.6
Sick leave	1.1	0.5	0.3
Group life	0.3	0.1	*
Nursing home	−1.0	−0.1	*
Retirement pension	−7.0	−0.5	−0.1
Taxes on earnings[a]	−1.1	−0.1	*
Differences in total net costs[a,c]	9.3	1.9	0.8
Life expectancy at age 20 (days)	−300	−22	−4

NOTE: * indicates figure is less than 0.05. All costs are external.
[a]Measured in thousands of 1986 dollars.
[b]Excludes maternity and well care.
[c](Sum of costs) minus taxes on wages. Due to rounding, categories may not sum to total.

TABLE 5—Sensitivity of Costs of Inactivity to Assumptions on Effects on Mortality (5 Per cent Discount Rate)

	Alternative Mortality Assumptions		
	Effect of Exercise is 1.5 More Years for Men	No Effect of Exercise on Mortality	HRA Model
Costs[a]			
Medical care[b]	1.8	2.0	1.2
Sick leave	0.5	0.5	0.5
Group life	0.1	0.0	0.4
Nursing home	−0.1	0.0	−0.3
Retirement pension	−0.5	0.0	−3.3
Taxes on earnings[a]	−0.1	0.0	−0.5
Total net costs[a,c]	1.9	2.5	−1.1

[a]Measured in thousands of 1986 dollars.
[b]Excludes maternity and well care.
[c](Sum of costs) minus taxes on earnings. Due to rounding, categories may not sum to total.

costs.

The discounted external lifetime costs, which are more relevant to policy, tell a similar story. At a 5 percent discount rate, increasing exercise saves $1,900 in lifetime costs. This represents our best estimate of the external costs of a sedentary life-style. By way of comparison, at a discount rate of 5 percent lifetime external costs of relatively inactive persons are greater than those of smokers ($1,000), but less than those of heavy drinkers ($4,600).*

Sensitivity Analysis

We analyzed the influence of alternate assumptions concerning the effects of exercise on mortality (Table 5). For convenience, the first column of Table 5 repeats the results from the second column of Table 4. The second column gives the costs assuming exercise has no effect on mortality; that is, the standard United States life table is used for both groups. Only medical costs and sick leave differ, and these are both higher for sedentary people than they would be if these same people were active.

The last column is based on the unaltered HRA results which show enormous beneficial effects of exercise on mortality (see Appendix). The mortality changes strongly affect pensions and use of nursing home care. If lack of exercise has a large effect on mortality, sedentary individuals pay much more in Social Security taxes than they live to collect. Even when discounted at 5 percent, these late-life effects mean that those who do not exercise subsidize those who exercise: the external costs of inactivity decrease as the assumed beneficial effect of exercise on mortality increases.

Next, we investigated the sensitivity of our results to other assumptions and data sets, while holding the effect on life expectancy at its middle value of 1.5 years for men. For comparison, the first column in Table 6 repeats the first column of Table 5. First, to test the sensitivity to data source, we used parameters based entirely on NHIS data (for the young as well as the old) (column 2). The major change was that the estimated effects of exercise on medical costs were much larger, which caused the external costs of a sedentary life-style to double. However, the sick leave and nursing home estimates are consistent with the HIE results.

Second, to test the sensitivity to how we estimated the health effects of exercise, we contrasted sedentary individuals with current exercisers, rather than with the sedentary people made hypothetically active (third column of Table 6). Because inactive people have other characteristics which are associated with lower medical use (for example they have less education, on average), they spend less on medical services than active people, despite their inactivity. The negative signs in the third column show that currently inactive people are not subsidized by exercisers; the first column however shows that inactivity is subsidized. In other words, if the people who are inactive started to exercise, they would pay current exercisers $1,900 more than the $700 they currently pay. The main reason inactive people do not impose higher external costs is that they die earlier.

The base case included all medical use, but some of the difference in medical use between inactive persons and "active inactive" persons may be unrelated to exercise. As a sensitivity test, we therefore narrowed the definition of medical services to those thought to be related to exercise and other habits. The fourth column gives the resulting medical costs when this restriction is imposed. (Here, we returned to the inactive group made hypothetically active.) Medical and total external costs fall by roughly a factor of two.

Fourth, many people become less active later in life. We therefore tried to estimate the costs that such a switch to inactivity impose on others. In doing so, we assumed that these people were similar to moderate or heavy exercisers up to age 50 and then became sedentary. Because the differences

TABLE 6—Sensitivity of Costs to Assumptions at 5 Per Cent Discount Rate

	Alternative Assumptions					
	Inactive Individuals if Active	All NHIS Data	Active Individuals	Narrow Definition of Medical Costs[b]	Switch at age 50	Total Costs
Costs[a]						
Medical costs[c]	1.8	4.4	−0.4	0.9	1.1	2.6
Sick leave	0.5	0.3	0.4	0.4	0.1	1.3
Group life	0.1	0.1	0.1	0.1	*	0.1
Nursing home	−0.1	−0.1	−0.1	−0.1	−0.1	−0.1
Retirement pension	−0.5	−0.5	−0.6	−0.5	−0.4	−0.5
Taxes on earnings[a]	−0.1	−0.1	0.1	−0.1	*	−0.7[d]
Total net costs[a,c]	1.9	4.3	−0.7	0.8	0.8	?

NOTE: * Indicates figure is less than 0.05. All costs are external, except last column.
[a]Measured in thousands of dollars.
[b]Only costs in diagnoses that are "probably related" to exercise.
[c](Sum of costs) minus taxes on earnings. Due to rounding, categories may not sum to total.
[d]Earnings, not taxes on earnings.

due to a sedentary life-style start late in life, sick leave and medical costs effects are diluted by discounting (column 5). At a 5 percent discount rate, external costs are $800 larger than if this group exercised throughout their lifetime.

Finally, the last column in Table 6 gives data on total costs rather than external costs; that is, it includes the portion borne by the person. Total costs are by definition higher than external costs in areas in which people pay part of the costs themselves. The costs shown in the last column amount to $4,000 but that figure does not count the cost to the sedentary person and his or her family of premature death and disability. Although this latter cost is probably larger than any of the costs shown, it is hard to quantify, so we have left the total cost figure in the lower right hand corner as a question.

Discussion

Because there has been no direct trial of the lifetime costs of not exercising, we have had to use several observational studies to estimate the components of costs. The main uncertainty in our overall estimate of lifetime costs comes from the validity of the assumptions, rather than statistical noise. Therefore, we have included extensive sensitivity analyses that show how our computed costs vary with the assumptions.

If sedentary individuals were more active they would live longer and reduce the costs they impose on others. The reduced costs come from lower covered medical and work loss costs associated with an active life-style. The higher taxes paid by active individuals over their longer lifetimes more than offset the additional pension payments received by active people as a result of their increased life expectancy. The effect of exercise is appreciable; for example, the external costs of a sedentary life-style are almost double the external costs of smoking.

In contrast to drunken drivers, the costs that sedentary people impose on others are indirect; sedentary people do not drive their armchairs into innocent bystanders. Most of these external costs are captured in the model, but we assumed there were no effects of inactivity on early retirement. In a study of those obtaining permanent disability Social Security benefits in 1975, 38 percent of retiring workers ages 55–64 had cardiovascular disease as their primary diagnosis.[19] Assuming the year 1975 was typical, we can estimate that three per cent of those ages 55–64 retire early because of cardiovascular disease, to which lack of exercise can contribute. Early retirement has large external costs associated with it, both because it decreases taxes paid and increases pension and disability insurance payments. If a sedentary life-style increased the probability of early retirement, our estimates of the external costs of a sedentary life-style are too low.

One of the largest uncertainties in our calculations concerns the degree of causality in the link between exercise and health status. Although we controlled for physical limitation and several other differences between exercisers and non-exercisers, we cannot be certain that the association we have found between inactivity and health-related costs is completely causal. In principle, this issue could be settled by a randomized experiment of the effects of exercise promotion, but size, cost, ethics, and long-term adherence problems make such an experiment impractical. The issue of causality remains to plague any observational attempt to estimate external costs of inactivity.

The association between exercise and mortality raises the same chicken-and-egg issue. Many epidemiological studies have shown dramatic differences in future heart disease between those who are sedentary and those who exercise.[5,6,20] These studies have controlled for age and other health habits. Still, it is hard to rule out the possibility that heavy exercisers are inherently more healthy than sedentary people in unmeasured ways, so that inactive persons who take up exercise will not enjoy the gains in life expectancy that models based on those studies would predict. If exercising has no effect on mortality and morbidity, then there are no benefits to making sedentary people more active. Because of this uncertainty, the revised HRA model 3.0 says only that exercise is probably good for you, and does not attempt to quantify its life-extending benefits.

Nevertheless, individuals contemplating a switch to a more active life-style will probably benefit from doing so.[5] According to Paffenbarger and Hyde,[4] the time spent walking is just returned in later life (undiscounted), and joggers who burn calories twice as fast as walkers can get double their exercise time back in increased life expectancy. However, the time given up to exercise is now, while the life extending benefits of exercise accrue in the future. Hence, if people care more about time now than time later,[11] they should engage in exercise that they either enjoy or has other benefits for them (e.g., increasing their feeling of well-being, physical appearance, etc.).

From a policy perspective, our results provide an economic rationale both for government spending on recreational facilities that encourage a more active life style, as well as for employer spending on worksite wellness programs that encourage exercise. About one-sixth of the population is sedentary but not physically limited. Whether subsidies to encourage exercise would induce enough additional exercise among that group to justify their costs is an issue we must leave to others; if our estimate of $1,900 in benefits per active person is approximately correct, a relatively small additional percentage of exercising would be enough to justify a subsidy, especially because the majority who already exercise would also derive benefits from such facilities and programs.

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APPENDIX
Further Details of Methods

1. Basis of Life Tables

We constructed an abridged life table using 1980 data on national mortality and the 1984 Centers for Disease Control's version of the Health Risk Appraisal (HRA) program developed by Robbins and Hall.[21] The HRA was used to adjust survival rates to reflect differences in the relative risks of dying between people who exercise and those who lead more sedentary lives.

Unfortunately, some problems in the 1984 HRA model's treatment of exercise lead to implausibly large estimates of the effects of exercise. For example, according to computations that we performed using the HRA model, men who switch from a sedentary life style to exercise will live nine years longer, almost double the computed effects of not smoking! This estimate is implausibly large for two reasons: 1) the HRA fails to account for the fact that better health may lead one to exercise, rather than vice versa; and 2) the adjustment made to the HRA model's exercise risk multipliers for age is inadequate, which makes the computed effects of exercise for older males too large. These two problems were compounded by the fact that our measures of exercise do not match well with the HRA categories. For example, because we had no information on past exercise habits, we had to assume that current patterns represent lifelong patterns. Yet exercise habits are not particularly stable.

Given these problems, we decided to estimate the external costs of exercise in three ways. As an upper bound on mortality effects, we used the unaltered results of the HRA model. For a lower bound, we assumed that exercise had no effect on mortality whatsoever, so that external costs arose only from differences in medical and sick-leave costs for sedentary and active people. For our best estimate, we uniformly altered the age-specific relative mortality risks of not exercising in the HRA model, so that male exercisers lived one and a half years more than male nonexercisers. This value was derived from several recent studies.[5,6,10] We did not adjust the mortality values for women because very little is known quantitatively about the effects of exercise on women and because the effects of exercise in the HRA model were not nearly as large as for men.

2. Assumptions on Collective Financing

To simplify the calculation of how much active and inactive people pay annually to finance these programs, we assume that each pays a given proportion of earnings, where the proportion is just enough to finance these programs (10.3 percent of earnings, at a 5 percent rate of discount). The assumption of proportionality is approximately correct for publicly financed goods and services,[22] and is also probably correct for privately financed pensions. Private premiums for health insurance are regressive, but this is somewhat offset by the benefits of Social Security payments and Medicaid financing for nursing home services being progressive.

3. Medical Care Services and Insurance Coverage

We examined the use of medical care services (excluding dental care) by those who exercise varying amounts. In the case of the HIE (RAND Health Insurance Experiment), medical expenditure measures were based on claims filed during the experiment for periods up to five years, and the individual's pre-experimental insurance coverage was used to estimate external costs. In the case of the NHIS (National Health Interview Survey), medical measures were outpatient visits and hospital admissions in the preceding year. Visits and admissions were converted to expenditures using average prices from the National Medical Care Expenditure Survey (NMCES),[23,24] and data on insurance coverage of the non-elderly from NMCES were used to estimate the portion of medical costs that were external.

In our analysis, we examined differences in medical use between those who engaged in regular exercise and those who did not. Because some of the observed differences in medical care use may not be caused by the amount one exercises, we made three corrections:

- We excluded maternity services because it seemed implausible that exercise was causally related to the decision to have children. (We also excluded well care because it is generally not covered by insurance.)
- To increase the chances that the difference in use of medical services is caused by differences in the

amount one exercises, we limited one set of analyses to diagnoses thought to be directly related to bad habits.
- We controlled for a set of confounding variables, as described in the text.

4. Work Loss

In the HIE, work loss includes time lost from work due to one's own or someone else's illness, including time for visits to a physician. A half-day or more missed from work was counted as a work-loss day. Because we do not know the reason for the work loss, we cannot restrict our measure to sick days due to illnesses known to be related to exercise. Instead, we estimated the difference in total work loss among persons with different levels of exercise, controlling for health habits and other covariates.

Wage data from the CPS were used to convert work loss days into costs. To obtain a figure for external costs, we used an estimate that employers pay 38 percent of the cost of work loss.[25]

5. Comparison of the HIE and NHIS Data Sets

We compared the estimated effects of exercise on medical costs and work loss on the non-elderly in the NHIS and the HIE data. Although there were few statistically significant differences, in some cases there were appreciable differences in the estimated coefficients. The differences in the magnitude of the estimated responses to exercise between the HIE and NHIS prompted us to perform the cost analysis two ways: 1) we used data from the HIE for those under age 60 and the NHIS for those 60 and older; 2) we used data exclusively from the NHIS (i.e., for all age groups).

6. Wages and Pensions

Data on wages come from the Current Population Survey.[16] Data on pensions, public and private, by age, sex, and education class also come from the Current Population Survey; they have been adjusted for the estimated 21 percent underreporting factor in this survey.[17] (We include private pensions as collectively financed because most are defined benefit plans rather than defined contribution plans and because defined contribution plans generally convert to annuities whose payout is independent of exercise status.) We assume survivors in each age, sex, and education class get the average pension for that class, i.e., we assume amount of exercise has no effect on age at retirement.

Cancer Prevention Fellowship Program Invites Applications

The Division of Cancer Prevention and Control (DCPC) is accepting applications for the Cancer Prevention Fellowship Program (CPFP). The purpose of the program is to attract individuals from a multiplicity of health science disciplines into the field of cancer prevention and control. The program provides for:

- Participation in the DCPC Cancer Prevention and Control Academic Course;
- Working at NCI directly within individual preceptors on cancer prevention and control projects;
- Brief field assignments in cancer prevention and control programs at other institutions.

Funding permitting, ten Fellows will be accepted for up to three years of training, beginning July 1, 1990. Benefits include selected relocation and travel expenses, paid federal holidays, and participatory health insurance.

Eligibility for the program is restricted to the following categories:

- MD or DO degree from a US, territorial, or Canadian Medical School, or
- PhD, DrPH, or other doctoral degree in a related discipline (epidemiology, biostatistics, and the biomedical, nutritional, public health, or behavioral sciences), or
- Foreign education, comparable to that received in accredited US, territorial or Canadian institutions. Foreign medical graduates must have current ECFMG/FMGEMS certification and appropriate experience, e.g., one year residency in a training program approved by the Accreditation Council for Graduate Medical Education.
- US citizenship or resident alien eligible for citizenship within four years.

The deadline for receipt of applications is September 8, 1989. To receive further details and application packet, sent a *postcard* with your name and home address to: Douglas L. Weed, MD, MPH, PhD, Cancer Prevention Fellowship Program, Division of Cancer Prevention and Control, National Cancer Institute, Executive Plaza South, T-41, Bethesda, MD 20892. Further inquiries: Mrs. Barbara Redding (301) 496-8640 or 496-8641.

Multivariate Prediction of Total and Cardiovascular Mortality in an Obese Polynesian Population

DOUGLAS E. CREWS, PHD

Abstract: The effects of body weight and blood pressure on the risk of total mortality and mortality from cardiovascular diseases (CVD) were examined in a prospective sample of 5,866 adult residents of American Samoa, a Polynesian population noted for exhibiting high levels of obesity. Data collected during 1975–76 were linked to mortality records from 1976 through 1981. In logistic regression models which did not include blood pressure, percent of desirable weight was an important risk factor for mortality from CVD, but it was not an important risk factor when diastolic blood pressure was included in the model. Percent of desirable weight was not related to mortality from all causes combined in either Samoan men or women. Age and diastolic blood pressure were predictors of total and CVD mortality in men and women.

These results, in an obese population, suggest that body weight and obesity are not independently related to excess mortality in the very obese, although they may associate with high blood pressure. These results also suggest that relations between physiological characteristics and mortality may vary with cultural, genetic, or other factors not examined in this study. (*Am J Public Health* 1989; 79:982–986.)

Introduction

The Medico-Actuarial Mortality Investigation of 1912[1] showed an increasing probability of death with increasing levels of body weight or blood pressure. Evidence that elevated blood pressure leads to early mortality, particularly mortality from cardiovascular diseases (CVD), has continued to accumulate,[2–6] but the influence of overweight and obesity on total and CVD mortality has been questioned.[6–12] At a recent National Institutes of Health (NIH) Consensus Development Conference on the Health Implications of Obesity,[13] most participants supported the view that obesity was predictive of early mortality, particularly from the vascular diseases,[14–16] but at least two papers reported some results that were counter to prevailing opinions.[9,17]

Age, weight, blood pressure, and fatness are generally accepted as risk factors for early mortality in western industrialized populations.[2–5,12,18–24] Although most prospective studies have assessed these relations in Western or White populations, data from the Honolulu Heart Program suggested that such associations also occur in Japanese Americans.[25]

Several genetically and culturally distinct populations reside on islands of the South Pacific. Of these, American Samoans are particularly noted for their heavy body weights, overt obesity, and hypertension.[26–30] Furthermore, this mainly Polynesian population exhibits high mortality rates from CVD and diabetes mellitus.[28,31]

As part of a longitudinal research program to examine the health consequences of modernization on American Samoans,[32] a prospective investigation was conducted of the relations between age, body weight, and blood pressure, and six-year mortality from all causes and from cardiovascular diseases (CVD). Multivariate analyses are reported here; univariate associations have been reported elsewhere.[33]

Methods

Study Sample

American Samoa is a group of high volcanic islands, lying in the South Pacific about 15° below the equator and just east of the international dateline. Tutuila, the main island, upon which over 90 percent of the population of American Samoa resides, comprises a total land area of less than 76 square miles. In 1974, 67 percent of adult males and 37 percent of adult females were employed and earning weekly wages,[34] with the remainder generally engaged in traditional subsistence and household activities.

Since 1962, the LBJ Tropical Medical Center has been a major provider of health care in the islands and has housed the registration office for vital statistics. The medical center also houses the Public Health Department which maintains nurses in clinics throughout the islands. In general, the nurses are familiar with the health problems of all households in the villages for which they are responsible.[35] Since almost all births and deaths are reported to either the outlying public health clinics or the medical center, there is little opportunity for incomplete ascertainment of vital events.[36]

Risk Factors

During 1975 and 1976, members of the Pennsylvania State University Samoan Studies Project (SSP)[32] and the Department of Public Health of American Samoa collected data on age, height, weight, systolic (SBP) and diastolic blood pressures (DBP) on 2,704 men and 3,162 women ages 20 to 90 years residing in 52 different American Samoan villages located on Tutuila. Percent of desirable weight was computed as weight (kg)/Metropolitan Life Insurance desirable weight specific for sex and height.[7,8] Percent of desirable weight was interpreted as an index of obesity.

All risk factors were measured according to standardized techniques.[30,37] However, some digit preference in blood pressure measurements has been previously noted in the Department of Public Health sample.[38] Methods and techniques of data collection and descriptions of the data have been presented elsewhere.[28,30,32,38,39]

Characteristics of the Sample

Summary statistics for survivors, all decedents, and those who died of CVD are presented in Table 1. The average age of the women sampled was 43 years, while the men averaged 44 years. In general men and women weighed the same; but men were approximately 10 cm taller. Thus, body mass index (BMI, computed as weight (kg)/height(m)^2) was higher in women. For comparison to published height and weight data from United States samples,[7] this sample of

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TABLE 1—Means (X̄) and Standard Deviations (SD) of Cardiovascular Risk Factors of American Samoans at Baseline Examination in 1976 by Vital Status on January 1, 1982

	Men						Women					
	Survivors (N = 2591)		Total Decedents (N = 113)		CVD[a] Decedents (N = 42)		Survivors (N = 3104)		Total Decedents (N = 58)		CVD Decedents (N = 21)	
Risk Factors	X	SD	X	SD	X	SD	X	SD	X	SD	X	SD
Age (years)	44.2	12.0	58.8	12.6	63.1	12.4	43.4	12.4	57.4	12.6	56.1	14.2
Weight (kgs)	86.9	18.1	87.0	21.0	92.4	22.4	85.2	18.9	82.8	27.7	81.4	25.2
Height (cms)	170.0	7.3	169.4	7.4	170.2	8.2	160.0	6.8	158.5	6.2	156.3	5.1
Body Mass Index (wt/ht^2)	30.0	5.7	30.2	6.5	31.8	7.1	33.3	7.0	32.7	9.7	33.1	9.3
Percent of Desirable Weight	135.4	26.0	136.0	29.5	143.0	32.2	160.5	33.5	157.3	47.6	158.9	46.6
Systolic Blood Pressure (mmHg)	129.1	18.6	140.5	25.0	151.3	27.2	127.7	22.8	155.1	37.0	159.1	47.1
Diastolic Blood Pressure (mmHg)	84.4	12.9	88.7	16.0	93.9	18.7	81.6	14.3	92.4	18.5	97.4	23.9

[a]CVD = Cardiovascular diseases

Samoan women averaged 5'3" in height and weighed 187.0 lbs, while the Samoan men averaged 5'7" and weighed 191.8 lbs. Among Samoan women and men these weights were, respectively, 35–41 percent and 27 percent above the midpoints of optimal weights for height recommended in the 1959 Metropolitan Life Insurance Height/Weight Tables and the Gerontology Research Center 1985 Weight for Height tables.[7] Neither SBP nor DBP showed sex-related differences.

The total sample represented approximately 59 percent of the Tutuila population aged 20 years or more in 1974. The age distribution of the sample was not significantly different from that of the 1974 census population.[38] The presence of cardiovascular and other diseases was not determined at baseline; nor were adequate data available to control for smoking habits, both possible confounders of the weight-mortality relationship.

Mortality Data

In general, death certification in American Samoa follows the pattern commonly used in the United States.[36] During January 1982 all official death records registered with the Department of Vital Statistics at the LBJ Tropical Medical Center between January 1, 1976 and December 31, 1981 were microfilmed. In 1972, death registration in American Samoa was reported to be nearly 100 percent complete.[36] Even if a few migrants were missed, previous research suggests they probably differed little in mean blood pressures (BP) from the residents.[29,30]

Underlying causes of death were categorized following protocols described in the 9th revision of the International Classification of Diseases, Injuries and Causes of Death (ICD-9).[40] The mortality data were then linked to the previously described survey data using the name, age, sex, and village of residence data from the death certificates. One hundred seventy-one decedents, 113 men and 58 women, were identified. Of these, 42 men and 21 women had died of CVD, (ICD-9 codes 401-459). The 2,591 men and 3,104 women for whom no death record was found provided the survivor sample. Previous estimates suggested 88 percent complete ascertainment of six-year decedents.[33]

Statistical Methods

Unadjusted associations of the risk factors with mortality were reviewed. The relative risk of mortality by categories of Metropolitan Life Insurance desirable weights,[5] and by quartiles of DBP, adjusted for age and sex, were estimated using multiple logistic regression.[41,42] Because few deaths occurred in the lowest two quartiles of DBP, they were combined to provide the base category for the analysis of mortality. Models with multiple predictor variables were used to estimate independent effects of percent of desirable weight and DBP on mortality controlling for age. Body weight unadjusted for height was not included in the logistic models. BMI and SBP were dropped because percent of desirable weight and DBP provided better fits to the mortality data. Risk functions including two factor interactions and quadratic effects were estimated. None of the age or weight by blood pressure interactions or quadratic terms was statistically significant; these exploratory analyses have been presented elsewhere.[28,39]

Four terms from each logistic model were presented, the estimated coefficient relating each risk factor to the endpoint along with its standard error (SE), the estimated relative risk (RR) of dying associated with a 10-year increase in age, a 10 mmHg increase in DBP, or a 10 percent increase in desirable weight, and a 95% confidence interval (95% CI) for the estimated RR.

Results

At their initial examinations in 1976, men and women who subsequently died within six years were 14.6 and 14.0 years older than those who survived (Table 1), for CVD deaths men were 18.9 years older, but women were only 12.7 years older than survivors. Absolute body weight, BMI, and percent of desirable weight were slightly but not significantly higher among men who subsequently died or died of CVD than they were among six-year survivors (Table 1). Samoan women showed higher levels of obesity and lower ages at death than did men. Both SBP and DBP were higher at baseline among Samoans who died than among those who survived six years. Men and women who survived showed similar blood pressure levels. However, among both total and CVD decedents, women showed higher SBP and DBP than men. Age-adjustment did not substantially alter any of these relations.

The relative risks (RR) of mortality from CVD and from all causes, adjusted for age and sex, according to desirable

weight levels are shown in Table 2. The pattern of estimated relative risks suggests a U-shaped function for mortality from cardiovascular diseases and all causes, with a graded relationship of increasing risk at each level of desirable weight above the 90–109 percent class. For the markedly overweight, percent desirable weight greater than or equal 160 percent, the RR of mortality from CVD was three times that of the 90–109 percent group. Despite this, when the 54 individuals with percent of desirable weight below 90 percent were deleted and the logistic function was used to test for a trend of increasing mortality with higher ranks of desirable weight, controlling for age and sex, no statistically significant trend was evident by the chi-square test.

Table 2 also presents relative risks according to quartiles of DBP. The RR of mortality from CVD and all causes combined increased with higher quartiles of DBP.

To distinguish the effects of DBP and weight seen in Table 2, logistic regression models were fitted for total and CVD mortality by sex. These models included age, DBP, and percent of desirable weight. The 54 individuals for whom percent of desirable weight was less than 90 percent were excluded from these analyses. Age and DBP were independent predictors of mortality from all causes in both Samoan men and women, but percent of desirable weight was not an independent predictor (Table 3); an increase of 10 percentage points in desirable weight was associated with a relative risk of only 1.04. Among Samoan women, a 10-year increase in age yielded a relative risk of death of 2.08, somewhat less than for men. Conversely, a 10 mmHg increase in DBP was associated with a 46 percent increase in the risk of death among Samoan women, twice the increase associated with DBP among Samoan men.

Age and DBP significantly increased the risk of six-year mortality from CVD among both Samoan men and women. Percent of desirable weight showed a weak positive association with risk in men and a weak negative association in women when the effect of DBP was also considered. Among Samoan men, age and DBP more strongly influenced the risk of mortality from CVD than the risk of mortality from all causes. Similarly, the coefficient for percent of desirable weight was higher for CVD than for death from all causes, among men. Among Samoan women age influenced mortality from CVD less than it did mortality from all causes combined, but DBP more strongly associated with the risk of CVD death.

Discussion

At a recent NIH Consensus Development Conference on the Health Implications of Obesity,[13] Garrison and Castelli[15] reported that 26-year mortality was lowest among Framingham men classified into the 100–110 percent group of the 1959 Build and Blood Pressure Study's desirable weight ranges. In this study of Samoans the lowest risk of mortality, especially from CVD, occurred in a similar category of relative weight, those 90–109 percent of Metropolitan desirable weights. The trend toward higher risks of mortality from all causes and CVD at both higher and lower percents of desirable weight suggested the now familiar U-shaped risk function. With individuals below 90 percent of desirable weight excluded from the analysis a trend of increasing mortality across the upper four categories of desirable weight was evident but not statistically significant in this sample. Estimates from a multiple logistic risk function suggest that elevated body weight and obesity are risk factors for total and CVD mortality in the extremely obese, but that this is secondary to the association of obesity with high blood pressure.

A strength of the study was the large and complete sample of the adult population of Tutuila. Therefore, the weak effect of body weight on mortality, adjusting for blood pressure, is probably not an artifact of the sampling design.

The average BMI in this Samoan sample was significantly above levels related to increased mortality in United States samples,[33,45] as was the average percent of desirable weight. Thus, the average Samoan man or woman in this study may have been above some critical threshold, after which the cardiovascular sequelae of obesity no longer increased. Although this suggestion may appear to contradict the finding that above 160 percent of desirable weight, excess CVD mortality was observed (Table 2), the effects of BP were not accounted for and both sexes were combined due to small numbers of observed deaths. Furthermore, at such extreme levels of body weight, mechanical and other unmeasured concomitants of extreme body weight may have contributed to the observed association.

A second possibility is that a six-year follow-up was too short a time period for the cardiovascular sequelae of obesity to be fully expressed. In the Framingham study the relation between relative body weight and mortality was weaker during the first 12 years of follow-up than during the second 12-year period,[45] however other studies with follow-up

TABLE 2—Relative Risk of Death, Adjusted for Age and Sex, from all Causes and Cardiovascular Diseases among American Samoans According to 1959 Metropolitan Life Insurance Desirable Weights and Quartiles of Diastolic Blood Pressure (DBP), Determined Using Logistic Regression[a]

Percent of Desirable Weight	N	All Causes			Cardiovascular Diseases		
		Cases	RR[b]	95% CI[c]	Cases	RR	95% CI
<90	54	7	2.070	.778, 5.508	2	1.450	.272, 7.732
90–109	504	23	1.000		7	1.000	
110–129	1213	37	1.033	.594, 1.799	12	1.200	.458, 3.141
130–159	2688	70	1.158	.695, 1.930	25	1.641	.681, 3.955
160+	1407	34	1.425	.788, 2.576	17	3.110	1.190, 8.130
Diastolic Blood Pressure Quartile mmHg							
1 and 2 <84	3305	56	1.000		17	1.000	
3 84–90	1388	52	2.040	1.366, 3.047	20	2.366	1.211, 4.623
4 >90	1173	63	2.535	1.690, 3.804	26	2.780	1.430, 5.402

[a]Percent of desirable weight and quartile of DBP were determined separately by sex, age was controlled and relative risks determined in separate logistic models for desirable weight and DBP.
[b]RR = relative risk of death compared to the 90–109% category of desirable weight or quartiles one and two of DBP.
[c]95% confidence interval for the estimate of relative risk.

TABLE 3—Multiple Logistic Risk Functions for Six-year Mortality from all Causes and Cardiovascular Diseases Based on Baseline Measurements of Risk Factors: American Samoa 1976–81

	All Causes								Cardiovascular Diseases							
	Men				Women				Men				Women			
Risk Factor	Coeff.	SE	RR[a]	95% CI[b]	Coeff.	SE	RR	95% CI	Coeff.	SE	RR	95% CI	Coeff.	SE	RR	95% CI
Age (years)	.0919	.0088	2.51	2.11–2.98	.0733	.0116	2.08	1.66–2.61	.1295	.0157	3.65	2.68–4.97	.0568	.0188	1.76	1.22–2.55
Diastolic Blood Pressure (mmHg)	.0202	.0075	1.22	1.06–1.42	.0380	.0083	1.46	1.24–1.72	.0365	.0108	1.44	1.16–1.78	.0520	.0114	1.68	1.35–2.10
Percent of Desirable Weight	.0036	.0042	1.04	0.96–1.21	−.0042	.0045	0.96	0.88–1.05	.0104	.0062	1.11	0.98–1.26	−.0069	.0071	0.93	0.81–1.07

[a]RR = relative risk associated with an increase of 10 years in age, 10 mmHg in DBP, or 10 percent in desirable weight.
[b]95% confidence interval for the estimate of relative risk.

lengths similar to this one have reported significant associations between obesity and mortality.[46]

Third, it is possible that the obese in Samoa were younger than in other populations, and as a consequence this population has not yet experienced the cardiovascular sequelae of their obesity. Negative correlations which existed between body weight or BMI and age[28,33] suggest this may be the case. An earlier analysis implied that overweight may influence mortality outcomes among younger Samoans.[47] Thus, as younger cohorts age the cardiovascular sequelae of their obesity may become more pronounced.

Several risk factors not included in this study, such as smoking and alcohol consumption or diet, may be confounding factors in the obesity mortality relationship in Samoans. Recently, Lew[16] reiterated the need to control such confounding factors before determining the effect of weight on mortality. Additional data to control for smoking habits might have helped explain part of the U-shaped relation between obesity and mortality and may have increased the relative risks reported for percent of desirable weight in the logistic analysis. However, the relative risks among Samoan women, who were a good deal more obese than men and seldom smoked cigarettes, were well below unity, suggesting that addition of a smoking variable may have had little effect.

Many studies have reported no association between body weight and mortality. Keys' review of 13 prospective studies found that only one showed a definite univariate relationship between overweight and CHD.[43] Similarly, review of 16 prospective studies of obesity and mortality led Andres to conclude that obesity does not influence total mortality the way one would expect.[33] What the present study indicates, as Barrett-Connor suggested based on a number of previous studies of well-defined populations,[9] is that it is often difficult to find a relation between body weight and increased mortality in populations unselected for risk factor status. This is particularly true in studies of nonwestern, nonindustrialized populations, and holds for mortality from all causes and CVD.[9,11,17,32,33,43,44] This suggests that at least two additional factors previously suggested by Stallones,[17] different research methods and true differences between the study populations, may have influenced these results. These two factors may also have interacted to accentuate differences between studies in Samoa and the United States.

The sample used in this study was composed mainly of Polynesian residents of American Samoa. They have a common genetic and cultural heritage, which is not the same as that of populations studied in the United States. It is possible that population heterogeneity has been a major difference between studies reporting a body weight-mortality relationship and those reporting none. Thus, differences in research samples, rather than designs may explain some observed differences in risk factor associations.

There also may be true differences between populations in the relation between body weight and mortality. These differences may stem from genetic factors relating to energy metabolism, storage, or expenditure and may have both an evolutionary and ecological explanation. It may also be that specific mechanisms have evolved in certain populations, such as the Polynesians of the South Pacific, so that they tolerate larger amounts of body fat without experiencing cardiovascular problems. In this context it may be rewarding to think in terms of a "thrifty genotype" as suggested by Neel.[48]

Given the extreme body sizes of present-day Samoans, it is easy to suggest that this population has an exaggerated ability to process food energy into fat. Such a physiological mechanism may also have conferred some degree of protection from the cardiovascular responses commonly observed in populations with greater genetic diversity and a different evolutionary history. This is similar to Alexander's suggestion that extreme obesity may carry some protective effect against the development of coronary heart disease.[49]

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APHA Epidemiology Section to Sponsor 'Late-Breaker' Poster Session

The Epidemiology Section of the American Public Health Association has issued a call for abstracts of papers suitable to feature in their "Late-Breaker" Poster Session, to be held during the upcoming APHA meeting in Chicago. The session will take place on Wednesday, October 25, 12:30–2:00pm. Work completed in 1989 is eligible for consideration.

Abstracts must be received by September 11, 1989 to be considered. Students and recent graduates are particularly encouraged to submit their work for this session. Abstracts should be less than 200 words; any format is acceptable. Enclose a self-addressed postpaid envelope to expedite the response from the program planner. Contact:

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Necrotizing Enterocolitis Mortality in the United States, 1979–85

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Abstract: The Multiple Cause of Death Mortality data from the National Center for Health Statistics were analyzed to describe epidemiologic characteristics and trends in deaths related to necrotizing enterocolitis (NEC) among infants in the United States from 1979 to 1985. The average annual mortality rate (multiple cause mortality) for NEC was 13.1 deaths per 100,000 live births. NEC annual mortality rates decreased significantly during the study period for White and Black infants, male and female infants, and infants in the Northeast, North Central, and South regions. Black infants were approximately three times more likely to die from NEC than White infants, and the NEC infant mortality rate was highest in the South. (*Am J Public Health* 1989; 79:987–989.)

Introduction

Necrotizing enterocolitis (NEC) is a serious gastrointestinal tract disorder in newborns,[1–3] affecting predominantly premature or low-birthweight infants. The overall incidence is approximately 1 to 3 cases per 1,000 live births[3–8] and 1.0 percent to 7.7 percent of all admissions to neonatal intensive care units.[1,9–12] Death associated with NEC results from intestinal perforation with sepsis, disseminated intravascular coagulation, and short bowel syndrome following intestinal resection.[1,13] Case series suggest that NEC mortality rates may be as high as 56 percent; however, rates have varied by study[1,8,14] and may have decreased over time.[14–16]

Before 1979, the International Classification of Diseases (ICD) did not provide unique coding for NEC as a cause of death. With the adoption of the ninth revision of the International Classification of Diseases in 1979, a cause-of-death code (777.5) was made available specifically for NEC. This provision allowed for the differentiation of NEC from gastroenteritis and colitis as a cause of death in death certificate data. In this study we analyzed NEC-related deaths in the United States from 1979 to 1985 for epidemiologic characteristics and mortality trends.

Methods

Multiple Cause of Death Mortality data for the United States from 1979 to 1985 were obtained from the National Center for Health Statistics (NCHS).[17] Cause-of-death classifications by NCHS from death certification data used the Ninth Revision of the International Classification of Diseases (ICD-9).[18,19] NEC deaths were considered those in which ICD-9 code 777.5 appeared in any one of 20 record axis codes,[17,18,20] i.e., deaths in which NEC was recorded as any one of the multiple causes of death. The record axis coding is a method which links causes of death into categories for certification by NCHS, and lists up to 20 causes of death for an individual.[20] The record axis data classification is comparable to the underlying cause classification, since it uses the same rules and category definitions.[17,20] However, the multiple cause of death data allows for mortality statistics based on a number of different causes reported on the death certificate.[20,21] Israel, *et al*, discussed the usefulness and limitations of multiple cause data in detail.[21]

The multiple cause of death data enabled us to examine NEC as an important factor in death. The NEC deaths included infants who died within the first year of life, plus two infants with unknown age. Detailed natality public use tapes from NCHS were used to determine the number of live births in each subgroup annually from 1979 through 1985.[22,23] All rates in this study represent multiple cause mortality rates, unless otherwise specified.

Annual NEC infant mortality rates based on multiple cause were calculated as the number of NEC deaths per 100,000 live births. The overall rate (multiple cause) for the study period (1979–1985) was calculated as the total NEC deaths per 100,000 live births during the study period. The overall and annual mortality rates were calculated for the United States and by race, sex, and region. Relative risks (RR) with 95 percent confidence intervals (CI) were calculated to compare specific mortality rates.[24] Linear regression analysis by the unweighted least squares method was used to describe the relationship of mortality rates to time.[25]

To examine racial differences and trends, race was categorized as White, Black, and other races. Geographic differences and trends were analyzed by standard census regions—Northeast, North Central, West, and South.

Results

From 1979 to 1985, 3,327 NEC deaths in infants were reported in United States multiple cause data, an average mortality rate of 13.1 deaths/100,000 live births per year (Table 1). Of these deaths, 79.6 percent were reported as the underlying cause of death, yielding an average underlying cause of death rate of 10.4 deaths/100,000 live births per year. During this period, NEC represented 1.1 percent of all infant mortality. Eighty-seven percent of NEC deaths occurred within the first month of life. The annual NEC mortality rate for the United States decreased significantly over time, from

TABLE 1—NEC Mortality Rate by Region and Race, United States, 1979–85

Region	Total (N)*	Race		Black/White RR (95% CI)
		White (N)	Black (N)	
Northeast	13.3 (628)	10.4 (402)	29.8 (218)	2.9 (2.4,3.4)
North Central	10.9 (707)	8.0 (441)	31.1 (258)	3.9 (3.4,4.5)
South	16.4 (1433)	10.3 (662)	35.0 (761)	3.4 (3.1,3.7)
West	10.0 (559)	9.0 (414)	26.0 (104)	2.9 (3.2,3.7)
United States	13.1 (3327)	9.4 (1919)	32.4 (1341)	3.4 (3.2,3.7)

*Includes other races (N = 67).

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14.5 deaths/100,000 in 1979 to 10.2 deaths/100,000 live births in 1985 (regression coefficient (b) and standard error = −0.75 (.12), R^2 = .88), a 29.7 percent overall decrease (Figure 1).

Overall, the NEC mortality rate for Blacks (32.4 deaths/100,000 live births) was significantly higher than for Whites (9.4 deaths/100,000 live births) (Table 1). The rate for other races (6.9 deaths/100,000 live births) was significantly lower than for Whites (RR = 0.7; 95% CI = 0.6, 0.9). For each year NEC mortality rates for Blacks were significantly greater than for Whites (Figure 1); however, relative risks remained similar over time. Annual rates declined significantly over time both for Blacks (28.3 percent) and Whites (31.1 percent) (Black, b = −1.58 (.54), R^2 = .56; White, b = −0.61 (.08), R^2 = .91).

The overall NEC mortality rate for males was higher than that for females (RR = 1.2; 95% CI = 1.1, 1.3) (Table 2). Annual rates declined significantly for both males (32.5 percent) and females (25.8 percent) (male, b = −0.74 (.22), R^2 = .70; female, b = −0.76 (.16), R^2 = .83). Both for males and females, the rate was significantly higher in Blacks than in Whites (Table 2).

The South had a significantly higher overall NEC mortality rate than the other three regions, and the West had a significantly lower rate than the Northeast (Table 1). The overall rate was significantly higher for Blacks than for Whites in all regions (Table 1). Among Blacks, the overall rate was highest in the South (35.0/100,000 live births); among Whites, the overall rate was highest in the Northeast (10.4/per 100,000 live births), and in the South (10.3/100,000 live births). Regression analysis showed a decrease in the rate from 1979 to 1985 in all regions (Northeast, b = −1.07 (.31), R^2 = .70; North Central, b = −0.75 (.15), R^2 = .84; South, b = −0.83 (.27), R^2 = .65; and West, b = −0.38 (.28), R^2 = .27) (Figure 2). The rate decreased over time both for Blacks and for Whites.

The overall rate for males was slightly higher than for females in all regions. Among males and females alike, the rate was lowest in the West, and highest in the South.* Among males, the rate was significantly higher in the South than in the North Central and West.

*Data available on request to author.

TABLE 2—NEC Mortality Rate by Sex and Race, United States, 1979–85

Sex	Total	Race		Black/White Relative Risk (95% CI)
		White	Black	
Male	14.3	10.6	34.5	3.3 (3.0,3.5)
Female	11.7	8.1	30.3	3.7 (3.4,4.1)
Total	13.1	9.4	32.4	3.4 (3.2,3.7)

Discussion

Our data corroborate the improvement in NEC mortality rates reported by others.[14–16] Potential causes for the decrease in NEC-related deaths include:

- a decrease in infants at risk for NEC because of improved prenatal care and a decrease in incidence of low birthweight and prematurity;
- a decrease in incidence of NEC among infants at risk because of prophylactic treatment (i.e., prophylactic antibiotics, parenteral nutrition, and/or IgA-IgG feeding);[26] and
- a decrease in mortality in infants with NEC because of earlier diagnosis, treatment, and surgical intervention.

The first two of these possibilities seem unlikely. Although there have been slow and recent declines in infant mortality,[27,28] including low-birthweight infant mortality, a parallel decrease in the proportion of low-birthweight infants has not been seen.[27] Also, although no data on the incidence of NEC are available during the study period, Finer, *et al*, and Kleigman, *et al*, reported an increase in incidence of NEC in the late 1970s.[9,10,28] Perhaps improvements in neonatal care, such as the institution of protocols discussed by Vermeylen, *et al*,[16] have allowed more premature infants to survive to develop NEC.

In our study, Black infants had more than a threefold higher risk of dying from NEC than White infants; this relative risk was consistent throughout the study period and by sex. Wilson, *et al*, reported a higher incidence of NEC in Black infants than White infants and a higher incidence in low-birthweight infants than in other infants.[4] After controlling for birthweight, however, the incidence and case fatality ratios did not differ between Blacks and Whites.[4] Kleinman, *et al*, also reported a higher incidence of low-birthweight

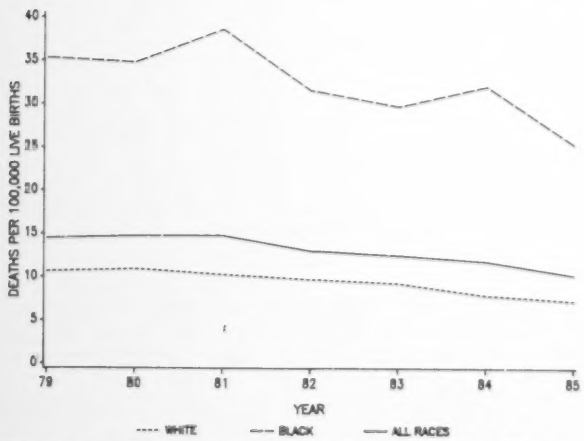

FIGURE 1—Annual NEC Mortality Rates by Race, United States, 1979–85

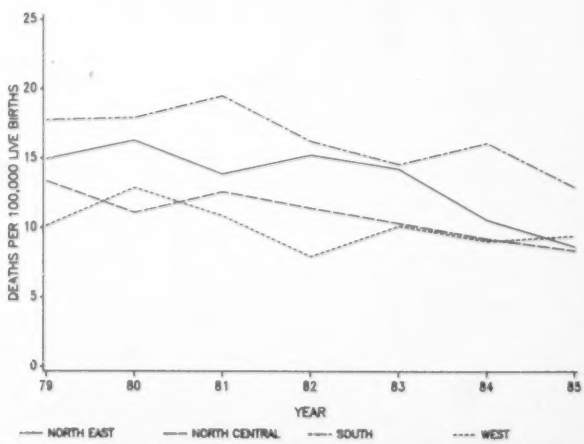

FIGURE 2—Annual NEC Mortality Rates by Region, United States, 1979–85

infants among Blacks.[29,30] During the 1970s, the incidence of low birthweight in Black infants in the United States was between two and three times higher than among Whites.[29,30] In 1980, the relative risk for low-birthweight Blacks was still greater than 2.0.[27] Therefore, the difference between Blacks and Whites in this study could reflect from an increased incidence of low-birthweight infants among Black infants. However, due to the limitation of the national mortality data, low-birthweight cannot be examined as an associated factor.

The sex differences that we observed in NEC rates have not been reported in previous case studies and are difficult to explain.[6,7,15] Male infants experience higher birthweight-specific infant mortality than female infants, regardless of race.[27] It is possible that the sex difference seen in NEC mortality can be attributed to the sex difference in overall infant mortality.

Infants in the South had a higher risk of NEC mortality than infants in the other three regions. In part this results from the higher proportion of Blacks in the South and the higher NEC mortality rate among Blacks. However, Blacks in the South had a higher mortality rate than Blacks in other regions. Blacks in the South do not have a higher incidence of low-birthweight births than Blacks in other regions.[31] The higher incidence of mortality due to NEC in southern Blacks may reflect less adequate postnatal care. However, there could be regional differences in the reporting of NEC. The use of standard cause-of-death classification does not assure strict comparability by region, which occurs only if all records of cause of death are reported with equal accuracy and completeness.[18] Since the reliability and accuracy is governed by the ability of the certifier, regional differences may reflect variations in reporting.[18,21,32]

Using national mortality data did not allow us to associate mortality with specific risk factors, such as birthweight, gestational age, clinical status at diagnosis, treatment, and maternal factors. NCHS will make available an annual national linked birth-infant death file for 1983–86.[33] Further studies will be able to expand upon our findings by examining the linked birth-infant death file for the United States with respect to the above risk factors, and should provide important information for the reduction of NEC mortality.

ACKNOWLEDGEMENTS

We thank Dr. Terry Chorba and Mr. Jack Smith for their helpful suggestions and review.

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Population-Based Study of Unintentional Injury Incidence and Impact during Childhood

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Abstract: We report a prospective study of medically treated unintentional injuries ascertained in an HMO population of children and adolescents 19 years of age and younger over a one-year period. The overall rate of medically treated injuries was 247 per 1,000, 147/1,000 for those treated in the clinic, and 100/1,000 for those treated in the emergency room (ER). Overall, 2.5 percent of patients were hospitalized. The highest rates of injury were for falls (60 per 1,000), recreational activities (57 per 1,000) and competitive sports (49 per 1,000). The site of care varied little by injury diagnosis or etiology, with the exception that sprains/strains, fractures, and injuries due to play and sports were more commonly treated in the clinic. The impact of trauma on activity was considerable, with 55.9 percent of injuries resulting in restricted activity, 10.6 percent in ≥2 days of school missed, and 6.7 percent in ≥2 days spent in bed. (*Am J Public Health* 1989; 79:990–994.)

Introduction

Injuries are the leading cause of death for children and young adults and a major cause of short- and long-term disability in the United States.[1] While advances in medical care have improved the outcome for some types of trauma, further significant decreases in morbidity and mortality will hinge on the development and implementation of effective prevention programs. As stated in the landmark report, *Injury in America*, "A prerequisite for the scientific study of injury is the acquisition of data on which to base priorities and research."[2]

Prior reports on non-fatal childhood injuries have focused almost exclusively on those injuries treated in the emergency room (ER) or hospital setting.[3–5] It is largely unknown, however, what proportion of all medically attended injuries are accounted for by injuries treated in the ER versus other sites; prior estimates place it at 40 percent for ER-treated injuries.[6] It is also not known in what way these injuries differ from those treated in a clinic or office setting, in terms of both etiology and type of injury.

Previous attempts to ascertain all medically attended injuries have significant limitations.[7–9] These include such problems as a recall bias and lack of important information on injury circumstances, etiology, treatment site, diagnosis, and severity.

The present study describes the incidence of unintentional injuries to children and adolescents 19 years of age and younger requiring medical care over a one-year period in a defined population, and compares the characteristics of injuries treated in the clinic to those treated in the emergency room setting.

Methods

A surveillance system was established to identify injuries treated in the clinic, ER, and hospital for children and adolescents 0–19 years of age who were members of the Northgate Medical Center of Group Health Cooperative (GHC) of Puget Sound. GHC is a large staff model health maintenance organization (HMO) serving approximately 330,000 members in the Puget Sound region.

Study Population

The group of children under surveillance during the study period October 11, 1985–October 10, 1986 were all enrolled as members of the Northgate GHC. This facility serves a population of 31,294 members, of whom 8,603 were 19 years of age or younger at the time of the study. As shown in Table 1, the study population is evenly distributed across age and sex groups. Nearly one-half of the study population heads-of-households had a college education, compared to 23 percent of the SMSA. The population also has a lower proportion of individuals in the lowest income category compared to the SMSA (14 percent vs 23 percent).

All GHC members are expected to and generally do receive medical care at their GHC clinic, emergency room, or hospital. In addition, audits of billing records for a 12-month period were done to identify patients with injuries treated outside the system and billed to GHC; these patients were included in the data presented. Finally, 7 percent of GHC enrollees have additional medical insurance. Injuries covered by this other insurance or paid out-of-pocket might be treated outside the GHC health care delivery system and would therefore be missed by the surveillance system established, resulting in a potential underestimate of injury rates.

Data Collection

Cases were identified through a daily examination of clinic and emergency room registration records by medical record technicians, using a specific protocol. Intentional injury cases, including child abuse, suicide attempts, and intentionally self-destructive actions were excluded. In addition, injuries due to complications of medical or surgical therapy, late effects of injuries, and illness due to food poisoning or contact dermatitis were also excluded. All other injuries falling within the International Classification of Diseases (ninth edition) codes 800–995 were included. An independent audit by study personnel of approximately 300 registration records found no injuries which had not been detected by this protocol. Age, sex, type of injury, and identifying data were abstracted from the record. Only the first visits for a new injury were included in the present analysis.

Injuries were sampled at 100 percent during the last eight months of the study; burns and admissions were sampled at

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TABLE 1—Socio-Demographic Characteristics of the Injured Population, Group Health Population, and the Local Standard Metropolitan Statistical Area (SMSA) (% in Each Category)

Demographics	SMSA[a] (n = 1, 607,469)	GHC (n = 332,000)	Northgate Medical Center (n = 31,924)	Injured Population (n = 2127)
Age (years)		% of 0–19 Population		
0–4	22.4	24.0	25.6	25.6
5–9	22.4	24.7	24.6	18.2
10–14	25.8	24.8	24.4	29.7
15–19	29.3	26.6	25.3	26.6
Sex				
Male	49	47	50	60
Female	51	53	49	40
Education of Head of Household				
High School or less	53	33	29	16[c]
College	23	49	48	43
Graduate School	24	18	23	34
Income[b]				
<$15,000	23	20	14	8[d]
$15,000–$34,999	37	47	53	41
≥$35,000	40	33	33	41
Marital Status				
Married/Living as Married	56	67	—	79[e]
Married	44	33		19

[a]SMSA 1980 Census.
[b]Income adjusted for a 43% increase in Consumer Price Index 1979–84.
[c]7% Unknown.
[d]10% Unknown.
[e]2% Unknown.

100 percent throughout the study period. The initial study (first four months) was undertaken as a thesis project by one of the authors (NC) and feasibility pilot with limited funding. Due to these circumstances, clinic and ER visits for injuries were randomly sampled at 25 percent and then 50 percent during the first four months of the study; sampling averaged 41 percent over this four-month period.

A detailed questionnaire about the circumstances of the injury and its consequences was mailed to the injured child's parent five to seven days after the initial visit to the emergency room or clinic. A reminder postcard was sent to all subjects not returning the questionnaire within one week. Non-responders were contacted by phone within four weeks of the medical visit and asked to complete the interview by phone. The overall response rate was 70.5 percent with 86.1 percent of the completed questionnaires done by mail and 31.9 percent by phone interview. Individuals responding by mail did not differ on the etiology of the injury, injury diagnosis, site of care, or household income from those in whom the questionnaire was completed by phone. These groups were therefore combined in the analysis.

Non-responders were compared to responders on age, sex, site of treatment, and type of injury. Non-responders did not differ from responders by sex distribution or site of treatment. The mean age of non-responders (11.70 years, SD = 5.36) was slightly greater than that of responders (10.50 years, SD = 5.30). Non-responders were also less likely to have more serious types of injuries such as fractures/dislocations (10.7 percent), lacerations (15.2 percent) and burns (2.7 percent) than responders (16.6 percent, 21.2 percent, and 4.2 percent, respectively).

Injuries were grouped into 14 different diagnostic categories according to ICD N codes, and into 12 different categories according to etiology or mechanism (Appendices A and B).

Rates of injuries were based on the 1,225 cases for which detailed questionnaire data were available. The total number of injuries in the study population was calculated by weighing for the sampling during the first portion of the study and for the non-response rate to the questionnaire in each sex and age category (<5 years, 5–9 years, 10–14 years, 15–19 years). Rates were calculated per 1,000 children using the number of registered members at the Northgate Medical Center in each age and sex group during the time period of the study.

Statistical testing was done using t-tests and Chi square as appropriate. Standard errors and confidence intervals were calculated based on the sampling and non-response rate.

Results

During the one-year study period, there was an estimated 2,127 injuries to the study population, for an overall incidence of 247 medically treated injuries per 1,000 children 0–19 years of age.

Age and Sex Variations in Rates and Sites of Treatment

As shown in Table 2, the majority of injuries, 1,266 or 59.5 percent, were treated in the clinic for an annual rate of 147 clinic-treated injuries per 1,000 children; the rate of ER-treated injuries was more than one-third lower at 100 per 1,000 (95 percent confidence interval of difference: 39.2, 54.8). There were 54 hospital admissions for an overall admission rate of 6/1,000 population or 2.5 percent of injuries. No child died from trauma in this population during the study period.

Consistent with previous studies, males had a higher rate of injuries than females, a difference which was greatest for the oldest age group. There were, however, no sex differences in the proportion treated in the clinic as compared to the ER. The rate of clinic-treated injuries was highest among 10–14 year olds, while 15–19 year olds had the highest rate of injuries treated in the ER. The proportion of injuries resulting

TABLE 2—Incidence of Injury by Age, Sex and Site of Treatment

	Rate per 1000											
	Clinic				Emergency Room				Total			
Age Group	Male	Female	Total	SE	Male	Female	Total	SE	Male	Female	Total	SE
0–4 years (n = 384)	113	74	94	10	99	61	80	8	212	135	174	13
5–9 years (n = 405)	127	122	125	10	77	56	66	7	204	178	191	12
10–14 years (n = 608)	217	185	201	15	141	90	116	10	358	275	317	18
15–19 years (n = 670)	218	119	171	14	167	102	137	11	385	221	308	18
Total (n = 2127)	169	124	147	6	122	77	100	5	291	201	247	8

NOTE: The population for each sex age combination is as follows:
0–4 includes 1112 males and 1093 females
5–9 includes 1077 males and 1041 females
10–14 includes 1074 males and 1029 females
15–19 includes 1147 males and 1030 females
SE = Standard Error

in admission was similar for all age groups with the exception of the 10–14 year olds where the proportion admitted was lower at 1.6 percent.

Injury Diagnosis and Site of Treatment

The most common injuries were sprains/strains, lacerations/punctures, contusions/abrasions, and fractures (Table 3). The site where medical care was received varied little with the type of injury. As might be expected, strains/sprains were predominantly treated in the clinic, and lacerations more commonly in the ER. However, fractures were treated more than twice as frequently in the clinic than the ER. Injuries such as contusions and abrasions were seen equally in both sites. The highest proportion of injuries requiring admission were internal injuries (100 percent), head injuries (13.3 percent), foreign bodies (14.3 percent), and fractures (7.5 percent).

Etiology of Injury and Site of Treatment

Injuries were classified into 12 mutually exclusive etiologic categories (Table 4). The highest injury rates were for falls, recreational activities, and competitive sports; these injuries were more commonly treated in the clinic than the ER. Injuries due to motor vehicle collisions, burns, and those in the other/unknown category accounted for the highest proportion hospitalized.

TABLE 3—Injury Diagnosis by Site of Care

		Rate per 100			
Injury Type	n	Clinic	ER	Total	SE
Sprain/Strain	503	40	18	58	4
Laceration/Puncture	426	21	28	49	3
Contusion/Abrasion	362	21	21	42	3
Fracture	335	27	12	39	3
Burn	70	5	3	8	.8
Head Injury	60	4	4	8	1
Dislocation	33	2	2	4	1
Foreign Body	26	2	1	3	.9
Bite	29	2	1	3	.7
Dental	6	0.3	0.4	0.7	.3
Poisoning	3	0.3	0.3	0.6	—
Internal	3	—	0.3	0.3	—
Drowning	1	—	0.1	0.1	—
Not Specified	220	22	4	26	3
Unknown	48	2	4	6	1
Total	2127	147	100	247	8

SE = Standard Error

TABLE 4—Etiology of Injury by Site of Treatment Rates per 1,000 Children 0–19 Years of Age

		Site			
Etiology	n	Clinic Rate	ER Rate	Total Rate	SE
Fall	513	33	27	60	3
Recreational Activities	494	36	21	57	3
Competitive Sports	422	33	16	49	3
Cutting Object	154	9	9	18	1
Motor Vehicle	127	8	7	15	2
Burn	70	4	4	8	.8
Other Person	61	4	4	8	1
Falling Object	31	2	2	4	.5
Poisoning	6	0.3	0.3	0.6	—
Drowning	1	—	0.1	0.1	—
Other/Unknown	246	18	11	29	2
Total	2127	147	100	247	8

SE = Standard Error

Injury Impact

Time in bed, time lost from school, and limitations of activity were used to measure the impact of the injury (Table 5). Patients seen in the ER were more likely to stay two or more days in bed and, where applicable, miss two or more days of school than were children treated in the clinic. Overall, 55.9 percent of patients had some limitation in activity for two days or more due to the injury; patients treated in the clinic were somewhat more likely to have two or more days of limited activity than were patients treated in the ER.

Impact of the injury also varied with age and etiology. Older children were more likely to spend two or more days in bed, miss two or more days of school, and have more limitations of activity than were younger children. Motor

TABLE 5—Morbidity Due to Injuries

	≥2 Days in Bed		≥2 School Days Missed		≤2 Days of Limited Activity	
	n	%	n	%	n	%
All Groups	139	6.7	195	10.6	1186	55.9
Site of Care						
Clinic	49	3.9	90	8.1	739	58.3
ER	90	10.8	105	14.5	447	52.5
Age (years)						
0–4	1	0.4	10	4.5	95	25.0
5–9	10	2.6	18	5.1	174	43.2
10–14	46	6.9	82	13.8	470	70.3
15–19	84	13.0	81	13.2	447	67.1
Etiology						
Motor Vehicle Accident	22	18.2	10	9.4	72	59.1
Sports	43	10.5	61	14.8	297	70.4
Play Falls	20	4.1	21	4.6	314	63.6
	22	4.4	44	10.9	235	46.2

vehicle injuries resulted in the highest proportion of children spending two or more days in bed, while injuries due to sports had the highest rate of children missing school or resulting in limitations of activity.

Location of Injury Occurrence

The site of injury occurrence varied with the age of the child, as shown in Figure 1. Overall, 28 percent of the injuries occurred in the child's home or yard, with the highest proportion of injuries in the home in young children (60.3 percent). In contrast, the majority of injuries in the oldest age group occurred in school or the schoolyard (25.6 percent), the street (14.2 percent) or the athletic field (30.2 percent).

Discussion

The rates of injuries treated in the ER and clinic found in this study are comparable to rates found in some prior reports,[6,7] while substantially different from those reported in other studies.[3,4] In particular, the rate of ER treated injuries was substantially lower (100/1,000 0–19 year olds) than that reported by the SCIPP[4] (216/1,000 0–19 year olds) and NEOTS[3] (205–255/1,000 0–14 year olds) studies. By contrast, the rate of medically treated injuries, 257 per 1,000 individuals 0–20 years of age, available from a special injury supplement to the 1975 National Health Interview Survey (NHIS)[6] is very similar to ours. One other study of an HMO population, performed 25 years ago by Manheimer, *et al*,[7] at a Kaiser Permanente clinic in Oakland, reported a rate of medically treated injuries identical to ours—246 per 1,000 0–15 year olds. The rate of hospitalized injuries in our study (6/1,000) was similar to the rate reported by the SCIPP study in their ER surveillance (7.7/1,000).

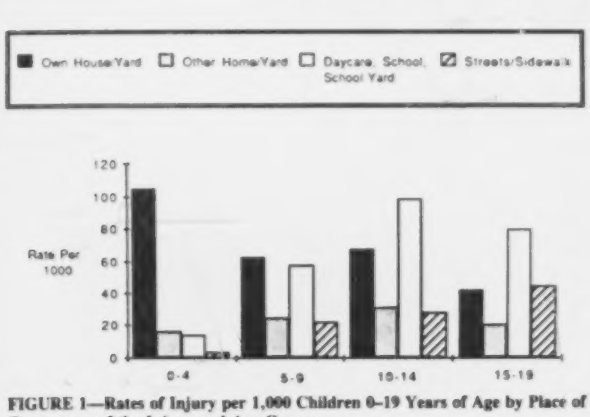

FIGURE 1—Rates of Injury per 1,000 Children 0–19 Years of Age by Place of Occurrence of the Injury and Age Group

These differences in rates may be a reflection of different socioeconomic groups in these studies, the form of medical care coverage available, or differences in care seeking behavior. Data from a number of studies clearly indicate that poor children have a higher rate of fatal injury than non-poor children, and probably a higher rate of certain types of non-fatal injuries as well.[10–12] Our study population was somewhat better educated and were less likely to be in the lowest income group than the surrounding SMSA; these differences may thus be partly responsible for the lower rate of ER-treated injuries reported.

Health care seeking patterns of members of HMOs might be different from the general population as well. At GHC, greater than 80 percent of members are specifically identified with a particular primary care physician. Under these circumstances it is reasonable to suppose that members will seek care at the lowest level (primary care) which can manage the problem, since this represents the specifically identified and most accessible care to the members. Thus, one might anticipate a shift of visits for injuries from the ER to the clinic. Although this likely occurred to some extent in our study population, the proportion of patients treated in the ER at GHC is similar to national data from the NHIS.[6] In that study, 46.7 percent of the medically treated injuries were seen in the ER compared to 40.2 percent in this study.

Our study has certain limitations which must be recognized. The incidence of injuries reported is probably an underestimate of the true rate because of non-GHC care covered by additional insurance or out-of-pocket expenditures by GHC members. As discussed earlier, 7 percent of members have the former; the extent of the latter is unknown.

The data are based on a sample (N = 1,225) out of a total of 2,127 injuries estimated to have occurred in the one-year period of surveillance. This is a reflection of the initial random sampling rates of 25 percent and later 50 percent in the first four months of the study. This procedure should not have introduced significant bias since the number of injuries were relatively large by type and etiological category, resulting in small standard errors and reasonable confidence intervals for most rates. In addition, the sampling did not appear to introduce a seasonality bias, since the proportion of injuries in each etiological category was similar during the first four months as compared to the subsequent eight months.

The data are based on a sample of patients, reflective of both the initial sampling scheme and the 70.5 percent response rate to the questionnaire. However, since the number of injuries was large the standard errors in most analyses were small, giving relatively narrow confidence intervals. Because of differences between non-responders and responders, the data underestimate injuries to younger children and less serious types of injuries such as contusions and abrasions.

The present study indicates that a large proportion of childhood injuries requiring medical care are treated in the clinic rather than in the ER setting. In our study, there were striking similarities in both etiology and diagnosis between those injuries treated in the clinic and those treated in the

emergency room. In addition, the study indicates that those injuries treated in the clinic result in a higher rate of limitations in activity. Thus, surveillance systems which collect data only on ER-treated injuries will miss injuries which are important (in terms of both frequency and impact) to the child and family. Injuries treated in both settings are important targets for prevention programs.

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APPENDIX A
Definition of Injury Diagnosis

ICD-9-n N Code	Injury Type	Definition
910–919	Contusion/Abrasion	Includes superficial injuries.
920–924		
870–872	Lacerations	Open wounds and traumatic amputations.
874–897		
	Fractures	Fractures to spine, trunk, extremities, face; excludes skull.
805–829		
830–839	Dislocations	All, including subluxations.
	Sprains/Strains	Sprains and strains of joints and muscles.
840–848		
940–949	Burns	All burns including scalds and flame.
800–804	Head	Fractures of skull; intracranial injuries; head injuries otherwise not specified.
850–854		
	Foreign Body	Foreign body entering through orifice; splinters.
930–939		
960–979	Poisoning	Poisoning by drugs, biologics, and non-medical substances.
980–989		
—	Asphyxiation/Choking	Asphyxiation by food or foreign object.
—	Drowning	Submersion.
—	Dental	Injuries to teeth.
	Internal	Internal injury of chest, abdomen, and pelvis.
860–869		
	Not Specified	Injury to body part not otherwise specified.
—		

APPENDIX B
Definition of Injury Etiology

	Injury Type	Definition
E 880–819	Falls	Excludes falls from motor vehicles, during competitive sport activities, and during recreational activities such as bicycling, skiing, jogging.
—	Play/Recreational	Includes bicycling, skating, jogging, skiing, everyday play.
—	Sports	Traditional competitive sports including baseball, basketball, football, soccer, hockey, etc.
E 810–819	Motor Vehicle	All injuries involving motor vehicles, including pedestrian- and bicycle-motor vehicle collisions.
E 820–825		
E 890–899	Burns	All burns including flame and scald.
E 924, 926		
E 850–869	Poisoning	Inhalation or ingestion of toxic substance, drugs other solid, liquid or gas.
E 910	Drowning	Submersion.
E 911–913	Choking, Aspiration, Suffocation, Asphyxiation	Mechanical suffocation, asphyxiation by food or other object.
E 916	Struck by Falling Object	—
E 919–920	Cutting Object	Injury by cutting or piercing object including machinery.
E 921	Struck by Other Person	Excludes child abuse.
—	Other	All other injuries.

A Case-Control Study of Pedestrian and Bicyclist Injuries in Childhood

I. Barry Pless, MD, Rene Verreault, PhD, and Sonia Tenina, MSc

Abstract: We identified children ages 0 to 14 years injured in traffic as pedestrians or bicyclists in Montreal, Canada. Two hundred children with injuries who received a score of 2 or more on the Maximum Abbreviated Injury Severity scale were considered as cases and compared with 400 uninjured children seen in the same hospitals for non-traumatic reasons. Systematic, blinded interviews and tests were conducted with parents to determine the role of a series of social, familial, personal, and behavioral characteristics. After adjustment for age, gender and socioeconomic area of residence, logistic regression analyses showed higher risks of injury to be related to fewer years of parents' education, a history of accident to a family member, an environment judged as unsafe, and poor parental supervision. Absence of physical health problems, fewer family preventive behaviors and reported lack of cautiousness were also related to a higher risk, whereas neither aggressivity nor behavioral disturbance, whether internalizing or externalizing, showed any such relation. These data suggest that the child's personality and behavior are weaker risk factors for pedestrian and bicyclist injuries than are family and neighborhood characteristics. (*Am J Public Health* 1989; 79:995–998.)

Introduction

Traffic accidents are the principal cause of mortality and morbidity among children.[1,2] Although in recent years much effort has been invested in descriptive studies of seat restraint use among child passengers,[3] and in evaluating programs designed to enhance use, much less is known about factors that place child pedestrians and bicyclists at risk for injuries in traffic. As a consequence, there are few well established strategies for the prevention of these injuries. Many of the most popular of these strategies rely on health education approaches, the efficacy of which appears to be distinctly limited.[4] Strong evidence has been presented to demonstrate that boys are more frequently injured in traffic than girls, and that living in lower socioeconomic areas is a risk factor.[5] However, assertions that behavioral factors[6,7] contribute to "accident-proneness"[8,9] or that family factors increase the risk of injury[10] are still open to question.

The present study describes the strength of a series of hypothesized risk factors for child pedestrian and bicyclist injuries using a case-control research design. Since many of these events result in injuries of minor severity and hence of little public health significance, the analyses reported here are restricted to cases of greater seriousness, as defined below.

Methods

All children (N = 1,600) under age 15 years injured in a traffic accident from July 1980 to December 1981 were identified through regular monitoring of 10 hospitals in the metropolitan area of Montreal and after a complete review of police accident reports.[11] This group includes virtually all children resident on the Island of Montreal who were treated in an emergency room (ER) for an injury from a passenger, pedestrian, or bicyclist accident during the 18-month study period. Their characteristics have been described in a previous report.[12]

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Cases in the present study were 200 children who were seen in the city's two children's hospitals with pedestrian or bicyclist injuries that received a score of 2 or more on the Maximum Abbreviated Injury Scale (MAIS).[13] The score increases with severity of injury from 1 to a maximum of 6, for death. An MAIS score of 2, which was chosen as the minimum severity for inclusion, corresponds to fractures, dislocations, or deep lacerations. The cases included 130 hospital admissions. Thirteen children who died from traffic injuries during the study period are not included.

Two groups of non-injured controls were selected from patients treated for medical problems in the same ER as the cases. The first, group 1, included 200 children matched by hospital, age (±2 years), sex and socioeconomic area (SEA) of residence (ranked high, low, and middle from census tract variables, e.g., father's occupation, family size and housing conditions).[14] The second, group 2, intended as a random sample, included 200 children but was only individually matched on hospital and frequency matched on age in the following categories: under age 5, 5–9, and 10–14 years. Medical problems that prompted ER visits by controls included, for the most part, colds, ear infections, and other minor complaints.

After each subject was identified, an interview with the mother was scheduled either at home or in the hospital depending on the condition of the child. The interview took place two to four weeks following the ER visit. Interviewers were intended to be unaware of the status of the subject, although in some cases this was not possible. The interview included the Child Behavior Checklist[15] and a structured questionnaire about sociodemographic and family characteristics, safety of the child's environment, extent of parent supervision, and the child's behavior and temperament characteristics.

All items in the questionnaire were summarized into a series of distinct hypothesized risk factors. Some factors were created as indices using several items that were first dichotomized and assigned a value of 0 or 1 according to the postulated direction of risk based on current knowledge. Indices were then calculated by summing this value and categorized into three or four groups. For example, a "family stress" index was created from the reported occurrence of 19 specific stressful events in the family during the six months preceding the ER visit (e.g., death of a family member, divorce, problems at work or at school, financial problems). A "preventive behavior" index (five items) related to the child's behavior when crossing streets, the use of automobile seat

restraints, and the parents' attitudes toward safety. An "environmental risk" index was created from four items describing the parent's assessment of traffic density and general safety of the immediate neighborhood. Finally, a "parental supervision" index (seven items) concerned safety related restrictions and cautions given to the child. In the latter, children under age 6 years were arbitrarily classified as being "at risk" with respect to the following: awareness of dangers when crossing between parked cars, understanding road safety rules, and having received formal safety instruction.

The scores on the Child Behavior Checklist serve to supplement a set of bi-polar personality descriptors included in the interview. Abnormal scores on this measure, which has been standardized and has established norms,[15] identify disturbed children similar to those receiving mental health services. Those with abnormal scores can be further categorized into those with disorders that are primarily "internalizing" (neurotic, withdrawn) and those that are "externalizing" (antisocial, aggressive). The results of the Child Behavior Checklist, completed by the parent, are expressed as normalized T scores, with scores greater than 63 representing those above the 90th percentile and hence abnormal.

Associations between the hypothesized determinants and the risk of pedestrian or bicyclist injury are presented as relative risks (RR). All analyses were performed using unmatched procedures after combining the two control groups.[16] This decision was reached after examining the results from preliminary analyses with each control group separately, using both matched and unmatched techniques, and finding little difference. All estimates of RR were obtained by unconditional logistic regression.[17] Logistic models included age (under 5, 5–9, 10–14 years), sex and SEA (low, middle and high) as confounding variables, because these were used as matching factors. For each analysis, the model included a single exposure variable, being assigned a reference value within each category, and all confounding variables. The ratio of the coefficient of the exposure variable to its standard error is presented with its corresponding p value and allows testing of the hypothesis that ln(RR) increases linearly across categories as a test for trends.

Results

Table 1 gives the distribution of age, gender, and SEA for cases and both groups of controls. Group 1 is similar to cases by age and gender but not quite as closely matched by socioeconomic area as intended. Group 2 also differs with respect to the sex ratio and thus reveals the expected increased risk of pedestrian and bicyclist injuries among boys. The number of cases peak in the 5–9 year age group.

Relative risks (RR) and 95% confidence intervals (CI) are presented in Table 2 for social and family determinants of injury after adjustment for age, sex, and socioeconomic area of residence. A high level of parents' education was strongly related to a lower risk of injury. Previous accidents of any family member requiring hospitalization as well as a low level of parental supervision and a high index of environmental risk were also associated with an increased risk of injury. Larger family size was related to increased risk, but this trend was less pronounced. Mother's employment and family stress were weakly and inversely related to risk.

Table 3 relates the hypothesized personal and behavioral determinants to risk of injuries. The presence of a physical problem—such as a chronic disability or impaired vision or hearing—was related to a reduced risk. Conversely, children with a low index of preventive behavior or a low level of cautiousness were at significantly higher risk. Nonsociability was associated with a slight increase in risk, whereas level of aggressivity showed no relation.

The presence of a behavioral problem, as assessed by the Child Behavior Checklist score, was also not related to the risk of injury (Table 3). Neither an internalizing nor externalizing behavior pattern was associated with the risk of traffic injuries.

All variables were again assessed in individual logistic regression analyses. Additional adjustment of the effect of each risk factor for all other established factors in a single regression model was performed but had little effect on the estimates of RR and therefore is not presented.* Finally, the relations of social, familial, personal, and behavioral factors to the risk of injury were examined within levels of age, gender and socioeconomic area. A previous accident was related to higher risk mainly among children being in low socioeconomic areas and the relation of a high family stress index and low risk was only observed among children in high socioeconomic areas. No other differences of consequence were observed over each level of age, gender and socioeconomic area.*

*Data available on request to author.

TABLE 1—Distribution of Cases and Controls by Age, Gender and Socioeconomic Area of Residence

Variables	Cases (n = 200) N (%)	Controls Group 1 (n = 200) N (%)	Controls Group 2 (n = 200) N (%)	Controls All Controls (n = 400) N (%)
Gender				
Boys	137 (68.5)	137 (68.5)	108 (54.0)	245 (61.2)
Girls	63 (31.5)	63 (31.5)	92 (46.0)	155 (38.7)
Age (years)				
1–4	28 (14.0)	24 (12.0)	35 (17.5)	59 (14.8)
5–9	111 (55.5)	106 (53.0)	102 (51.0)	208 (52.0)
10–14	61 (30.5)	70 (35.0)	63 (31.5)	133 (33.3)
Socioeconomic Area of Residence				
Low	78 (39.0)	67 (33.5)	69 (34.5)	136 (34.0)
Middle	62 (31.0)	72 (36.0)	71 (35.5)	143 (35.8)
High	60 (30.0)	61 (30.5)	60 (30.0)	121 (30.3)

TABLE 2—Social and Family Determinants of Pedestrian and Bicyclist Injuries

Risk Factors	No. Cases	No. Controls	R.R.*	95% CI
Mother's Education				
<7 years	43	73	1.0	
8–11	110	181	1.0	0.6, 1.5
12+	45	145	0.5	0.3, 0.8
		X trend = −2.78	p = 0.01	
Number Children in Family				
1	36	92	1.0	
2–3	125	249	1.3	0.8, 2.1
4+	39	59	1.8	1.0, 3.1
		X trend = 1.91	p = 0.06	
Mother Employed				
No	112	191	1.0	
Parttime	36	92	0.7	0.4, 1.1
Fulltime	52	117	0.8	0.5, 1.1
		X trend = −1.58	p = 0.11	
History of Accident in Family				
No	118	277	1.0	
Yes	82	123	1.6	1.1, 2.3
Family Stress Index				
Low	98	160	1.0	
Intermediate	59	129	0.8	0.5, 1.2
High	43	111	0.7	0.4, 1.0
		X trend = −1.98	p = 0.05	
Parental Supervision Index				
Good	23	71	1.0	
Intermediate	102	238	1.3	0.8, 2.2
Poor	75	91	2.6	1.5, 4.7
		X trend = 3.63	p < 0.001	
Environmental Risk Index				
Low	49	174	1.0	
Intermediate	101	174	2.0	1.4, 3.1
High	50	52	3.4	2.0, 5.6
		X trend = 4.81	p < 0.001	

*Adjusted for age, gender and socioeconomic area of residence.

TABLE 3—Personal Determinants of Pedestrian and Bicyclist Injuries

Risk Factors	No. Cases	No. Controls	R.R.*	95% CI
Physical Problem				
No	102	206	1.0	
Yes	56	206	0.4	0.3, 0.6
Preventive Behaviors				
High	55	152	1.0	
Intermediate	117	222	1.5	1.0, 2.2
Low	28	26	3.0	1.6, 5.6
		X trend = 3.37	p < 0.001	
Cautiousness				
0 (cautious)	83	207	1.0	
1	62	110	1.4	0.9, 2.1
2	35	56	1.5	0.9, 2.4
3 (not cautious)	20	27	1.7	0.9, 3.2
		X trend = 2.02	p = 0.04	
Aggressivity				
0–1 (not aggressive)	25	61	1.0	
2	46	85	1.4	0.8, 2.5
3	64	141	1.2	0.7, 2.0
4 (very aggressive)	65	113	1.4	0.8, 2.4
		X trend = 0.77	p = 0.44	
Sociability				
0–1 (sociable)	32	81	1.0	
2	42	91	1.2	0.7, 2.1
3	58	107	1.4	0.8, 2.4
4 (not sociable)	68	121	1.5	0.9, 2.5
		X trend = 1.57	p = 0.11	
Behavioral Determinants				
Behavior Problem (Sum T > 63)				
Normal	139	269	1.0	
Abnormal	54	120	0.9	0.6, 1.3
Internalizing				
Normal	148	286	1.0	
Abnormal	45	103	0.8	0.5, 1.2
Externalizing				
Normal	141	302	1.0	
Abnormal	52	87	1.2	0.8, 1.8

*Adjusted for age, gender and socioeconomic area of residence.

Discussion

These findings are in sharp contrast to those of some investigators who report maladjustment or disturbed behavior [6,7,18,19] but consistent with others.[20,21] Although there is a large literature on behavioral abnormalities of children who have been injured, and those who are "accident prone," few of the former studies report consistent behavioral abnormalities. In contrast, two reports suggest that activity level and/or aggression are significantly associated with the occurrence of subsequent injuries.[20,21] Such findings are difficult to compare, because the nature of the injuries are often different, as are the measures of behavior. However, to the extent that over-activity may be equivalent to lack of cautiousness, our findings agree with those of the latter investigators using longitudinal data.

It is also often difficult to interpret what behavioral relationships with injuries mean. For example, one study[7] suggests that behavioral characteristics may predict injuries differently among "normal" healthy children who are simply being adventuresome, extroverted and curious, compared to those with clinical behavioral disturbances leading to withdrawal and who may have a diminished ability to cope with environmental demands and the tasks needed to ensure safety. Our findings suggest that behavioral disturbance as such plays little role in the prediction of injuries, whereas traits such as lack of cautiousness may be of some importance.

It is not surprising that having a physical health problem appears to reduce risk of injury, but the wording of the question does not enable us to determine whether the association exists because these children are less exposed, more cautious, or more closely supervised.

Of greatest interest are the remaining risk factors. The first of these relates to the nearly three-fold increase in risk among children in the lowest category of the preventive behavior index. Although this may partly reflect post hoc reporting problems, it appears that children in families where parents admit engaging in many unsafe preventive behaviors—such as not using seat restraints in motor vehicles—truly have an increased likelihood of having a traffic injury.

The credibility of this interpretation is supported by the finding that when any member of the family has been seriously injured previously, the risk to the child is increased by about 50 percent. (This is not confounded with family size since the finding remains after adjustment for all other risk factors.) Parental supervision was also self-reported and hence subject to possible distortion, but not to social desirability bias. This strong association is consistent with a pattern of risk as a family issue rather than a personal one. This too is in keeping with other reports.[23,24]

Also compelling are the strong relations with reported safety of the environment where most of the children were injured. Here again, the question of reporting bias must be acknowledged. Parents whose children have been injured may understandably exaggerate the degree to which their neighborhood is dangerous. Against this, a number of other

reports show strong associations between injury rates and areas characterized by few parks, few bicycle paths, poorly designed intersections, and dense traffic.[25,26] And, perhaps most importantly, these may also be areas where driving behavior is most dangerous.[27,28]

In summary, although these results provide some evidence that the behavior of the child or parent plays a role in increasing the risk of traffic injuries among children, they provide a much more compelling argument for examining more closely the risks associated with environmental factors. In communities where urban planning has taken the safety of the child into account, specifically through measures aimed at separating traffic from bicyclists and pedestrians, there appear to be significantly fewer injuries.[29,30] Although few of these reports provide conclusive evidence of the effectiveness of these preventive strategies, it seems promising that those responsible for developing injury prevention strategies for a community to take full account of such features of the environment. The alternative is to continue to "blame the victim," in this case the child, by focusing on imprudent behavior, or to further emphasize the parent's obligation to provide better supervision. Either of these strategies, while not entirely misplaced, are much less likely to result in a significant reduction in traffic-related injuries in the short run than would programs aimed at changing the behavior of the driver or improving urban environments through traffic engineering.

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4th National Conference on Chronic Disease Prevention, Control

The Centers for Disease Control (CDC) and the Association of State and Territorial Health Officials (ASTHO) will cosponsor the Fourth National Conference on Chronic Disease Prevention and Control: "Implementing Effective Strategies," September 20–22, 1989, at the Bahia Hotel in San Diego, California. The conference is open to the public; there will be no registration fee.

The conference will emphasize the interactions among federal, state, and local health departments, voluntary health agencies, professional organizations and others, and will include plenary sessions that address legislation, surveillance, and year 2000 health objectives. Concurrent sessions will focus on: screening and quality assurance; evaluation, diabetes and obesity; childhood antecedents of chronic disease; emerging issues, programs, and controversies in chronic disease prevention and control; barriers to risk-factor modification among minorities; and nutrition and cancer. For additional information, contact Jack Friel, Center for Chronic Disease Prevention and Health Promotion, Mailstop A37, CDC, Atlanta, GA 30333; tel: (404) 639-2249 or FTS 236-2249.

Comprehensive Health Care for High-Risk Adolescents: An Evaluation Study

FELTON EARLS, MD, LEE N. ROBINS, PHD, ARLENE R. STIFFMAN, PHD, AND JACK POWELL, PHD

Abstract: This study was designed to evaluate the effectiveness of a large scale program to improve health care for high-risk adolescents. Seven clinics, funded to provide comprehensive primary care to adolescents, were compared to three non-funded clinics. The majority of the 2,788 adolescent patients sampled in these clinics were female (78 percent) and Black (71 percent). Each patient was initially interviewed at the time of a clinic visit and reinterviewed 12 months later; their medical records were systematically reviewed. As expected, the funded clinics detected and treated a wider range of medical and behavioral problems than the comparison clinics, a finding that was based on the self-reports of patients and confirmed by documentation in the medical records. However, improvements in life-style and in specific medical outcomes were not observed. To build on the limited success of this program, efforts are needed to encourage more males to receive care and to develop more specific interventions for patients in this age group. (*Am J Public Health* 1989; 79:999–1005).

Introduction

The problems adolescents take to doctors challenge conventional approaches to medical care. To promote a focused and intensive effort to improve the health care of young people, adolescent medicine[1–4] has been quick to identify the major causes of mortality and the most common morbidities characterizing this age group[5]; less certainty exists on knowing how effective interventions have been in improving the health of adolescents, however.[1–5] To a large extent this is because the approach to problems such as unprotected sexual intercourse, substance abuse, and violent behavior require behavioral and life-style changes that go beyond traditional medical practices.

Results of a broad-scale program to improve the health care of high-risk populations of adolescents are reported in this paper. The project, introduced as the Program to Consolidate Health Services for High-Risk Young People and sponsored by the Robert Wood Johnson Foundation, was started in 1981. Twenty medical schools were selected to receive funds to provide comprehensive services to youths at higher risk than the national average for the major problems characteristic of this age group, and by improving the training of physicians in adolescent medicine.[6]

The concept was based more on the capacity to provide comprehensive health services than on the organizational characteristics of the participating programs. For this reason the evaluation was guided by five objectives:

1) Do the funded clinics attract truly high-risk youths?
2) Do they detect their problems?
3) Are comprehensive services provided?
4) Are these services implemented?
5) Are health outcomes improved?

To address these questions, the clinics receiving grants from the Johnson Foundation were compared to a group of clinics that did not have a special focus on adolescents. Because of their location in inner-city areas, all of the clinics were in a position to serve high-risk youths.

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Methods

Research Design

The research design was that of a two-wave survey in which youths attending primary health clinics were interviewed twice at a 12-month interval. Of the 20 funded clinics, seven were chosen to participate in this evaluation study because they had a relatively large volume of patients in the adolescent age range from which to sample. The 20 clinics linked medical school training programs in adolescent medicine to a range of services based in hospitals, neighborhoods, and schools. However, each program was given considerable leverage in choosing how to develop a consolidated program. Of the seven programs studied, five had clinics in multiple settings including three with school-based clinics; those that operated from a single locus had established ties with other services; all seven had multidisciplinary staffs that included social workers, psychologists, special education counselors, midwives, and recreational counselors as well as physicians and nurses.

The seven funded programs (all associated with academic medical centers in Boston, MA; New Haven, CT; Indianapolis, IN; Chicago, IL; Jackson, MS; Dallas, TX; and Los Angeles, CA) were compared to clinics that were part of three non-specialized services (located in Buffalo, NY; St. Louis, MO; and New Orleans, LA). The comparison clinics did not offer special services to teenagers, nor did they have training programs in adolescent medicine. They were typically staffed only with physicians and nurses and had few administrative ties to other community-based clinics; none operated school-based clinics. The funded and non-funded clinics were similar in that they were all located in major public medical centers in their respective cities and were accessible to indigent patients. A brief description of the 10 programs, highlighting some of their similarities and differences, is provided in Appendix A.

Sampling

Between 150 and 360 youths were sampled from each clinic between November 1984 and May 1985. The age range of patients was restricted to 13–18 year olds. The sample represented a series of consecutive and unduplicated teenage patients appearing on days when the clinics were operating. The variation in the number sampled reflected the volume of patients in this age range attending the clinics during this six-month period, and in that sense the patients are representative of a typical day in the operation of primary health

clinics. Most were interviewed on the same day that they attended the clinics in an on-site space where confidentiality could be maintained and their medical care not compromised.

During the six-month field period of Wave 1, 3,102 patients became eligible to participate in the study, based on their age and presentation at one of the clinics. Of these, 2,788 were interviewed. Among the 314 patients who were identified but not interviewed, 115 refused; 199 were not interviewed either because they were over the age of 18, too ill, or could not be located during or following their visit. Eighty-seven percent of those interviewed during Wave 1 were reinterviewed 12 months later during Wave 2. Because of the time consumed in locating and reviewing medical records, only about 83 percent of youths completing both the Wave 1 and Wave 2 interviews had their records abstracted. Complete data are available for 72 percent of the original sample.

There were no substantial differences in the number of patients completing all aspects of the study in the three types of programs, as shown in Table 1.

Interviews and Record Reviews

Highly structured interview protocols covered the youths' reasons for attending the clinics, other health services used, their physical and mental health status, family background, school adjustment, peer relations, stressful events, social adjustment, social supports, health practices, and extracurricular activities. Sections of the interview dealing with depression, anxiety, alcohol/drug use and abuse, tobacco use, and physical symptoms were taken from the Diagnostic Interview Schedule (DIS).[7] A section on aggressive and antisocial behavior was taken from the Diagnostic Interview for Children and Adolescents (DICA).[8] The follow-up interview, completed 12 months after the initial interview, covered these same issues and assessed change in health status and satisfaction with services received. During this second wave, the medical records of the youths were reviewed and abstracted.

At the Wave 2 interview, (75 percent of which was done at home) particular treatment for problems that had been identified in the Wave 1 interview was also queried. These problems included the presenting problem for which the youth was seeking help as well as any others reported during the interview session, regardless of whether or not they were known to health personnel at the study clinic.

TABLE 1—Sample Sizes at Wave 1 and Wave 2

		Of Those Completing Wave 1			
	Total Number Interviewed at Wave 1	Completed Wave 2 Interview		Completed Wave 2 Interview and Had Medical Record Review	
Type of Clinic	N*	f	%	f	%
Funded NH	1320	1087	(82)	925	(70)
Funded S	616	568	(92)	452	(73)
Comparison	852	760	(89)	625	(73)
Total	2788	2415	(87)	2002	(72)

*This number represents 86%, 97%, 92%, and 90% of all eligible cases in the Funded NH, Funded S, and Comparison clinics and the total sample, respectively.
NH - Neighborhood/hospital-based
S - School-based

A structured protocol, designed to complement the interview, was used to abstract information from the medical records. The interviewing and record reviews were done by professional interviewers (Survey Research Associates of Baltimore) who were trained prior to each wave of interviews. An effort was made to collect records from all services used by patients even if they were not affiliated with the study programs.

Results

Sample Description

Females made up over three-fourths of the sample and Blacks represented 71 percent. Black females constituted the largest sex-race group and were the most uniformly distributed across all the sites, while White males represented the smallest and least well distributed group. Of the three types of clinics, the school-based clinics had the largest proportion of males (43 percent), followed by the funded neighborhood/hospital-based clinics (21 percent), and the comparison neighborhood/hospital-based clinics (12 percent).

This patient sample is unrepresentative of the general population of youths who lived in the urban areas where the clinics were located. In some cases, youths originally residing outside these areas were attracted to these clinics because they wished to avoid stigma or sanctions that might have accompanied help-seeking in their own communities. Even among those living in the same areas as the clinics, self-selection operated to yield a group in which females predominated and males not in school were underrepresented. In the neighborhood and hospital-based clinics, 19 percent of the youths were not in school, a rate which varied from 4 percent for Black males to 42 percent for White females. Fifty-seven percent of the youths' families were headed by persons with blue collar jobs, and 7 percent of their breadwinners were unemployed. Nine percent of the youths lived in circumstances in which there was no parent or parent surrogate present.

Presenting Problems and Risk Factors

Table 2 shows the frequency of the presenting problems reported at the initial interview. Over three-fourths of the youths were attending for problems unrelated to illness (i.e., pregnancy, birth control, check-up, and preventive care), a percentage that varied from 61 percent in the funded neighborhood/hospital-based clinics to over 90 percent in the comparison clinics. Those attending the funded clinics had higher rates of utilization for acute medical problems, contraceptive needs, mental health problems, sexually transmitted diseases, and chronic illnesses than those in the comparison clinics; and lower rates for pregnancy. Not unexpectedly, the school-based clinics attracted the largest proportions of youths requesting check-ups and other forms of preventive care. These differences among the three types of clinics remained after controlling for age, race, and sex.

Table 3 shows the percentage of individuals attending the three types of clinics who had ever experienced a variety of life-style and behavioral problems. A cumulative index was derived by summing the presence of each problem over the life-span prior to Wave 1 and its occurrence (or recurrence) in the interval between Wave 1 and Wave 2 interviews. The definition of each risk factor is given in Appendix B. As a measure of highest risk, the proportion of youths with three or more of these risk factors is included in Table 3. These indicators represent some of the major risks for injury and social problems in adolescents and, quite apart from their

TABLE 2—Percentage Distribution of Presenting Problems of Adolescents at Three Types of Clinics

Presenting problem	Funded NH (N = 1087) %	Funded S (N = 568) %	Comparison (N = 760) %	Total (N = 2415) %
Pregnancy*	19	7	53	29
Check-Up/Preventive	20	49	17	26
Acute Medical Problem	26	19	11	20
Contraceptive Needs	23	5	17	17
Psychiatric	7	6	<1	5
Chronic Illness	7	3	2	4
Sexually Transmitted Disease	6	1	<1	3

*Females only (Funded NH, N = 878; Funded S, N = 326; Comparison, N = 674).
NOTE: Percentages total more than 100 because youths could present with more than one problem.

TABLE 3—Life-style and Behavioral Problems Ever Experienced by Youths at Three Types of Clinics as Reported at Initial Interview

Life-style and Behavioral Problems	Funded NH (N = 1087) %	Funded S (N = 568) %	Comparison NH (N = 760) %	Total (N = 2415) %
Physical Fights	67	71	62	66
Contraceptive Needs	60	58	69	62
Pregnancy	34	16	64	39
Regular Tobacco Use	37	16	33	31
Hopelessness	37	21	27	30
Heavy Alcohol/Drug Use	31	18	20	25
Runaway	24	13	20	20
School Drop-Out	19	4	27	18
Arrests	23	11	18	19
Use of Weapons	14	10	13	13
Suicidal Behavior	12	6	8	9
3+ Problems	60	41	64	57

presenting problems, should be of keen concern to health providers.

The school-based clinics had the lowest proportion of high-risk youths (41 percent) and these youths had lower rates on all but one (fighting) of the 11 individual risk factors compared to youths in the other types of clinics. Engaging in physical fights and unprotected sexual activity (shown by the proportion with contraceptive needs) were the most common problems. Suicidal behavior was the least common.

Pattern of Utilization and Care

To examine the impact of the funded programs, 11 problems were selected that included both traditional medical conditions, such as asthma and sexually transmitted diseases (STD), and behavioral and life-style problems, such as those listed in Table 3. The list, based on problems designated by the youths in the initial interviews regardless of whether or not such problems constituted the reason for seeking care, conforms to the types of problems targeted by the funding agency in its charge to improve services for high-risk youths.

Table 4 illustrates the extent to which this range of health problems was recognized and treated in the three types of clinics. The most important contrast is between the funded neighborhood/hospital-based and the comparison clinics, since they were located in similar settings. The funded school-based clinics provide data to contrast with the neighborhood/hospital-based clinics. The Table shows that youths in the funded clinics were more likely to inform providers about their behavioral and life-style problems than youths in the comparisons clinics. On the other hand, informing the clinics about more clear-cut medical problems (contraceptive needs, asthma, injury and STD) was done with greater consistency and at higher rates among all clinics than they were for behavioral and life-style problems.

Informing a clinic about a problem constitutes a first step toward getting help. Table 5 shows the proportion of youths who actually received help among those who had told a clinic about their problems. In most cases, help was obtained on-site. But because having the capacity to refer patients to other services when needed was a key element of the effort to consolidate services, it is also included in the Table. Youths with asthma and conduct or interpersonal problems in the funded clinics were more likely to be treated on-site than youths in the comparison clinics. What is most striking, however, is that the numbers of youths available to be treated in the comparison clinics were greatly reduced because fewer had informed the health care providers about their life-style and behavioral problems. Differences in the rate of referral among the clinics were not striking.

Medical Record Documentation

While the research interview was designed to comprehensively assess the health and psychosocial status of the youths, providers were faced with the need to specifically address the patient's presenting problem. Within the limited and often pressured time frame of most patient encounters, they also had to conduct a general examination of their patients and record their findings in the medical record.

We compared the youths' health and psychosocial profiles based on data from the interview to that abstracted from the medical records by counting the number of problems identified in each document using a common list such as that

TABLE 4—Percentages of Youths Who Returned to the Clinic between Waves 1 and 2 and Informed Providers about Their Health Problems in Funded Neighborhood/Hospital Clinics (N = 1087), Funded School Clinics (N = 568), and Comparison Clinics (N = 760)

Health-Related Problem		N with Problem	Percent with Problem Who Returned and Informed Clinic about Their Problem
Contraceptive Needs	FNH	606	72
	FS	272	50
	COM	485	66
Physical Symptoms	FNH	742	44
	FS	293	39
	COM	535	36
Depressive Symptoms	FNH	378	21
	FS	98	22
	COM	195	6
Conduct Problems	FNH	274	11
	FS	72	21
	COM	123	2
Asthma	FNH	115	55
	FS	55	44
	COM	69	58
Quitting Tobacco	FNH	258	24
	FS	46	11
	COM	168	12
Interpersonal Problems	FNH	525	10
	FS	211	11
	COM	350	3
Anxiety Symptoms	FNH	235	22
	FS	54	22
	COM	136	10
Injury	FNH	204	22
	FS	90	34
	COM	93	30
Alcohol/Drug Abuse	FNH	152	25
	FS	44	25
	COM	56	9
Sexually Transmitted Disease	FNH	61	57
	FS	14	43
	COM	22	54

FNH - Funded neighborhood/hospital-based clinics
FS - Funded school-based clinics
COM - Comparison clinics

shown in Table 4. The findings are shown in Figure 1. As expected, the interviews provided more information than the medical records. The largest discrepancy occurred in the comparison clinics. While youths in the funded neighborhood/hospital-based clinics and the comparison clinics had the same mean number of risk factors based on the interview data, the number of risk factors documented in the records was significantly higher in the funded clinics. This supports the self-reports of the youths and indicates that funded clinics were clearly superior to the comparison clinics in detecting a broad range of health problems in their patients.

Outcome Measures

Two approaches were taken to detect differences in outcomes between youths in the funded and comparison clinics. One approach was aimed at detecting changes in the number of behavioral and life-style problems in the 12-month period preceding the Wave 1 and Wave 2 interviews, and the other examined the degree of recurrence or persistence of specific medical or psychiatric problems. No evidence was found favoring the funded clinics in either case. The proportion of youths experiencing three or more risk factors in the year preceding the Wave 2 interview remained the same as it had been in the 12 months preceding the Wave 1 interview in both the funded and the comparison clinics. Even when age, race, sex, and the number of problems present at Wave 1 were controlled, the risk status of the youths at Wave 2 was not diminished in the funded programs. This was true not only when the proportion of youths with multiple problems was examined, but also for a reduction in the average number of behavioral and life-style problems reported in the year following the initial interview. In fact, there was a small increase in the proportion of youths with multiple problems suggesting that the interventions were not able to overcome the expectable increase in risk associated with the aging of an adolescent sample.

A number of specific outcomes were examined next. The frequency of sexually transmitted diseases in the interval between Waves 1 and 2 among youths with a previous history of such infections was not reduced in the funded clinics compared to the nonfunded clinics. This was also true for having a second (or third) pregnancy. In fact, for most of the problems examined, we failed to find evidence of a reduction in risk among youths attending the funded clinics. Although the proportion with unmet contraceptive needs remained uniformly high in all the clinics, we found, unexpectedly, that the proportion with persisting depressive symptoms and heavy alcohol/drug use at Wave 2 was actually higher in youths in the funded neighborhood/hospital clinics than in the comparison programs, as shown in Table 6.

Discussion

The answer to the first question of this evaluation study—"Do the funded clinics attract truly high-risk youths?"—is affirmative. Nearly two-thirds of the youths attending the neighborhood and hospital-based clinics had multiple behavioral and life-style problems that placed them at high risk for the social morbidities characteristic of adolescence and young adulthood and for early death. However, the proportion of such youths in the funded clinics was no greater than that in the comparison clinics. All the clinics turned out to be particularly appropriate places to reach high-risk females, a finding supported by another recently completed study of an adolescent clinic in a suburban setting.[9] While the school-based clinics had a lower proportion of high-risk youths, they attracted more males than did the neighborhood/hospital-based clinics. All of the clinics, however, failed to attract males who were not in school. This remains a challenge since these males represent the highest risk group for intentional and unintentional injury in the community. One approach to attracting a larger number of high-risk males is through liaison with juvenile detention facilities. This was pursued by one funded neighborhood/hospital-based clinic which, not surprisingly, had the highest proportion of males attending these types of clinics.

Clinicians in adolescent medicine are faced with detecting and treating a spectrum of problems for which traditional pediatric training may not have thoroughly prepared them. In addition to traditional medical problems, such as sexually transmitted diseases, and pregnancy, their experience must also alert them to the importance of behavioral and life-style problems such as depression, delinquency, substance abuse, unprotected sexual intercourse, and cigarette smoking.[10] For this reason we were especially interested in the answer to the second question of the evaluation regarding the detection of a broad range of health and social problems in adolescent patients. The superiority of the funded clinics in recognizing

TABLE 5—Percentages of Youths Treated, Referred, or Receiving No Care among Those Who Informed Providers about Their Problems in Funded Neighborhood/Hospital Clinics (N = 1087), Funded School Clinics (N = 568), and Comparison Clinics (N = 760)

Health-Related Problem		Of Those Who Informed Providers N	Percent Treated On-Site	Percent Referred	Percent with No Care
Contraceptive Needs	FNH	437	94	0	6
	FS	135	89	2	9
	COM	319	89	7	4
Physical Symptoms	FNH	328	72	3	25
	FS	113	76	8	16
	COM	195	69	3	28
Depressive Symptoms	FNH	80	76	15	21
	FS	22	95	18	4
	COM	12	75	17	25
Conduct Problems	FNH	73	63	10	36
	FS	46	83	2	17
	COM	10	20	10	80
Asthma	FNH	63	59	10	40
	FS	24	42	8	54
	COM	40	27	0	73
Quitting Tobacco	FNH	61	46	—	54
	FS	5	60	—	40
	COM	20	45	—	55
Interpersonal Problems	FNH	55	80	0	20
	FS	24	79	0	21
	COM	10	50	0	50
Anxiety Symptoms	FNH	52	58	0	42
	FS	12	92	0	8
	COM	13	46	8	46
Injury	FNH	45	62	9	29
	FS	31	77	7	16
	COM	27	70	11	19
Alcohol/Drug Abuse	FNH	38	47	0	53
	FS	11	73	0	27
	COM	5	40	0	60
Sexually Transmitted Disease	FNH	35	97	0	3
	FS	6	100	0	0
	COM	12	75	25	0

FNH - Consolidated neighborhood/hospital-based clinics
FS - Consolidated school-based clinics
COM - Comparison clinics

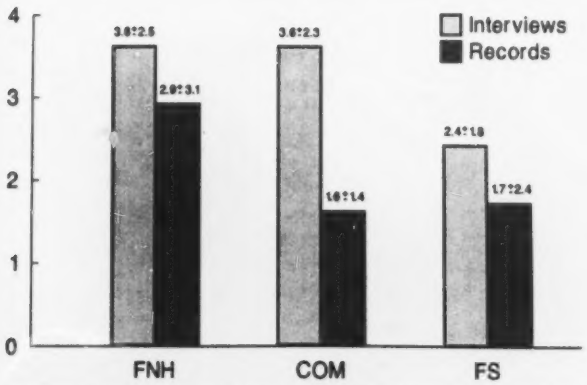

FIGURE 1—Comparison of the Mean Number of Adolescent Health Problems Reported on Interviews and Found in Medical Records
FNH = Funded Neighborhood/Hospital-Based Clinic
COM = Comparison Clinic
FS = Funded School-Based Clinic

behavioral and lifestyle problems was clearly shown by the higher proportion of youths reporting such problems to them and by their more extensive demonstration of such problems in their medical records.

No doubt the greater recognition of problems in the funded clinics was a function of their more diverse and specially trained staffs. Their success in this regard resulted in more youths receiving treatment and interventions being provided for a broader range of problems than in the comparison clinics. This answers the third question of the evaluation regarding the availability and implementation of comprehensive services.

It is worth emphasizing that, although the quality of care was not directly assessed, a higher quality of care in the funded clinics can be inferred from their more extensive documentation of problems.[11,12] Nevertheless, the funded

TABLE 6—Percentages of Youths with Persistent Problems from Wave 1 to Wave 2* in the Three Types of Clinics

Type of Clinic	Persistent Depressive Symptoms %	Unmet Contraceptive Needs %	Heavy Alcohol or Drug Use %
FNH (N = 1087)	23[a]	36	31[b]
FS (N = 568)	12	29	18
COM (N = 760)	14	33	21

*Persistence is defined as the presence of the problems in the 12 months prior to the Wave 1 interview and again in the 12-month interval between the Wave 1 and Wave 2 interview.
a) FNH - COM = 9% (95% CI = 5.5, 12.5)
b) FNH - COM = 10% (95% CI = 6.0, 14.0)

clinics were not able to alter the propensity of their patients to continue their adverse life-styles and to accumulate additional problems.

There are at least three reasons to explain why we failed to detect outcome differences:

- The follow-up period of one-year was a short interval over which to find differences in problems many of which have their bases in lifestyle orientations.
- More time may be needed to judge the impact of specialized primary health care for adolescents since pediatricians have only recently begun to develop skills in these areas. One goal of the Robert Wood Johnson Foundation was to inspire efforts to develop training programs and to promote the concept of interdisciplinary, consolidated health care for this age group. While the Foundation's initiative was clearly successful in supporting many new programs throughout the country, the long-term success of this field will depend on the extent to which stable support is found to maintain its focused agenda.
- A more sobering reason why we may not have detected outcome differences relates to the inherent nature of many of the problems characteristics of high-risk youths. Perhaps their difficulties are so deeply embedded in the economic and social contexts from which they come that improvements in their health status require more than medical clinics can provide. If this is the case, then the goals of treatment should relate more to the stabilization of risk conditions, than to their actual reduction.[13]

There are two limitations of the present study that deserve special mention. As in many evaluation studies, the research design for this one was developed after the programs had already been funded. This resulted in less comparability among the clinics in their organizational patterns and in the character and intensity of the interventions they provided. The other limitation relates to the possibility that biased reporting of health-related information occurred in the interviews. Efforts were made to minimize this problem by using highly structured and carefully validated interview methods, systematic training, and monitoring of interviewers, and by comparing information derived from the interviews with that included in medical records.

Overall, the evaluation demonstrates that the Johnson Foundation's objectives met with limited success. Continued progress in the identification of the behavioral and life-style problems possessed by these youths must be followed by finding ways to attract more males as patients and by devising interventions that, at a minimum, have the capacity to slow down the progression of life-style and behavioral problems that are typically occurring at their peak rates during the later adolescent years.

APPENDIX A

Ten Participating Clinics Included Seven with Adolescent Services (Funded Clinics) and Three with Non-specialized Services (Comparison Clinics)

Funded Clinics

1. A hospital-based adolescent clinic, a neighborhood-based adolescent clinic, and a school-based clinic were included in this program in New Haven, Connecticut. The multidisciplinary staff included an adolescent medicine specialist, fellows in adolescent medicine, a nurse-midwife, a social worker, and a psychologist.

2. A hospital-based adolescent clinic, a neighborhood-based clinic, and a school-based clinic were components of the program in Chicago, Illinois. The staff consisted of an adolescent medicine specialist, fellows in adolescent medicine, a nurse practitioner, a school nurse, a social worker, a guidance counselor, and a special education counselor.

3. Two high-school-based clinics, with neighborhood and hospital-based clinics serving as part of a consolidated network, composed the program in Jackson, Mississippi. The staff included an adolescent medicine specialist, fellows in adolescent medicine, a nurse, a social worker, and an educational counselor.

4. A hospital-based clinic and a neighborhood-based adolescent clinic made up this program in Boston, Massachusetts. The staff was composed of adolescent medicine specialists, fellows in adolescent medicine, a nurse, a psychologist, and a recreational counselor; the program also sponsors a school-based violence prevention program.

5. This program in Dallas, Texas consisted of a hospital-based adolescent clinic that placed emphasis on screening and diagnostic evaluation. It tied to an extensive referral network of neighborhood-based clinics throughout the city. The multidisciplinary staff included an adolescent medicine specialist, fellows in adolescent medicine, a nurse, a nurse practitioner, a social worker, and a consulting psychiatrist.

6. A neighborhood-based adolescent clinic and a closely associated hospital-based adolescent clinic constituted the two components of this program in Los Angeles, California. The staff consisted of a specialist in adolescent medicine, fellows in adolescent medicine, a nurse, a social worker, and an educational counselor.

7. The program in Indianapolis, Indiana was composed of several neighborhood-based adolescent clinics which were coordinated through a hospital-based adolescent medicine program. The multidisciplinary staff included an adolescent medicine specialist, fellows in adolescent medicine, nurses, and social workers.

Comparison Clinics

1. The clinics in New Orleans, Louisiana were all located in the largest public general hospital in the city. The hospital was part of a large academic medical center with no special clinics for adolescents. Patients were sampled through a walk-in clinic for all non-emergency cases, and through an obstetrical clinic. The staff included staff physicians, residents, and nurses only.

2. Publicly supported neighborhood health clinics with no special services for adolescents made up this service in St. Louis, Missouri. Adolescents were sampled through pediatric, dental, and obstetrical clinics. The staff included only staff physicians and nurses.

3. This site in Buffalo, New York was composed of general outpatient, emergency, and obstetrical clinics of a Children's Hospital and a hospital-based family practice center. The staff included physicians, nurses, and residents in pediatrics, obstetrics-gynecology, and family practice.

APPENDIX B

Risk Factors	Definitions
Pregnancy	Pregnant one or more times.
Use of Weapon	Threatened anyone with a knife or gun or any kind of weapon such as a pick, soda bottle, or something else.
Physical Fights	• Been in physical fights, *or* • Hurt someone badly in a fight, *or* • Taken things from other people by purse-snatching, holding them up, or threatening them.
Legal Problems	Been in trouble with the police or a juvenile officer.
Runaway	Run away overnight or longer from the place where living.
Contraceptive Needs	If had sex *and*: • respondent and partner do not use birth control most of the time, *or* • withdrawal is one type of birth control used.
Regular Tobacco Use	Smoked cigarettes, cigars, or pipes; chewed tobacco; or used snuff every day for a month or more.
Heavy Alcohol/Drug Use	• Drank four or more glasses/drinks of beer/wine/liquor at one time, *or* • Averaged drinking more than once a week or for short concentrated periods over the last year, *or* • Used a drug more than once a week or for short concentrated periods over the last year.

continued

APPENDIX B continued

Risk Factors	Definitions
Suicidal Behavior	Attempted suicide.
Hopelessness	• During a period of feeling blue or depressed (defined as: 1) two weeks or more during which respondent felt sad, blue, depressed, and lost interest and pleasure in things he/she cared about and enjoyed; *or* 2) two years or more when respondent felt sad or depressed almost all the time), work at school, at home, or on the job suffered, *or* • A period of two weeks or more when life felt hopeless or pointless.
School Drop Out	Not in school; highest grade completed was less than 12th.

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1988—A Landmark Year for State AIDS Legislation

A detailed legislative review of AIDS-related laws from the 1988 state legislative sessions show that out of 700 AIDS-specific bills introduced in 42 states, a total of 136 laws were passed in 36 states. Six state legislatures were not in session last year. The review, published by the AIDS Policy Center of the Intergovernmental Health Policy Project, found the 1988 legislation to be more comprehensive than in previous years. A series of omnibus laws address controversial and often conflicting policy issues: adding anti-discrimination provisions; amending public health control laws; developing informed consent and confidentiality provisions; and adding testing requirements for special populations or special situations.

Few laws concerning mandatory testing and reporting, contact tracing, quarantine and criminal sanctions were actually passed, although many were introduced. States tended to reject broad measures calling for mandatory testing of the general public and continued to emphasize voluntary testing and counseling, based on written, informed consent. Twenty-four states adopted measures requiring reporting of HIV-positive test results, including "identifiers" (usually a name and address) in at least some situations. However, most states also provide for anonymous testing. In four-fifths of the states, some form of anonymous testing is still available; only eight states have no option for anonymous testing.

During 1988, 27 states strengthened confidentiality provisions in existing communicable or sexually transmittable disease laws to cover test information or passed new laws to provide explicit confidentiality for HIV records.

Whether physicians have a responsibility or duty to inform the spouse or sexual partners of HIV-infected patients of their potential for exposure was addressed in laws in 10 states which now authorize a "discretion to warn."

A number of other major areas were addressed by the states, including voluntary partner notification, non-discrimination laws, restrictions on use of blood tests by insurance companies, and other related matters. For more information about the report, contact: AIDS Policy Center, Intergovernmental Health Policy Project, George Washington University, 2011 I Street, NW, Suite 200, Washington, DC 20006. Tel: 202/872-1445. The Center provides AIDS information and related research to legislators, policymakers and the media.

Psychoactive Drug Use in a General Population Sample, Sweden: Correlates with Perceived Health, Psychiatric Diagnoses, and Mortality in an Automated Record-Linkage Study

CHRISTER ALLGULANDER, MD

Abstract: Three person-based computer files were linked to provide a data-set of a random sample of 32,679 Swedes, drawn for interviews regarding perceived health, socioeconomic conditions, and psychoactive drug use. All diagnoses from inpatient psychiatric care in the sample during a 15-year period and the causes of death after the sampling point were combined with the interview responses. Among those admitted for inpatient psychiatric care, substance abuse was an infrequent diagnosis; the majority of schizophrenics and of those with an affective disorder appeared not to medicate regularly; survey non-responders had higher rates of mental disorders than responders. Drug use correlated with both subjective and objective measures of mental ill health. The rate of prescription drug abuse was low. Automated record-linkage is a feasible method to generate hypotheses about mental health in the general population. (*Am J Public Health* 1989; 79:1006–1010.)

Introduction

Psychiatric morbidity in general population samples has been surveyed to elucidate rates using lay and professional interviewers,[1,2] prognosis by computer-analyzed questionnaires,[3] and drug treatment of symptoms and diagnoses.[4–7] Correlates between psychoactive drug treatment and social and demographic conditions have also been investigated using pharmacy records,[8] health survey interviews,[9–13] personality inventory,[14] and insurance enrollment questionnaire.[15,16]

The purpose of this study was to describe the characteristics of persons reporting regular psychoactive drug use based on a large interview sample of the general Swedish population, and national psychiatric inpatient and cause-of-death registers, using automated record-linkage of three person-based computer files.

Methods

Four different random samples of the entire Swedish population 16–74 years old were interviewed in the Survey of Living Conditions in 1975, 1977, 1980, and 1981 (see Appendix for details). A total of 32,679 individuals were sampled, representing an average population of 5.8 million Swedes. Of these, 26,952 were interviewed; 5,727 (17.5 per cent) were not interviewed, due to refusals (11 per cent), being unavailable (2 per cent), ill (0.4 per cent), or institutionalized (0.3 per cent). We selected 20 appropriate questions for study (Table 1).

By means of the personal identification numbers, we matched the 32,679 individuals with 1.7 million cases of the national psychiatric inpatient case register for the period 1969–83. The ICD-8 diagnoses, slightly modified in Sweden, were grouped into ten nosological categories. A total of 1,765 sampled individuals had been discharged with psychiatric diagnoses during this 15-year period. Matching errors based on personal identification numbers, were estimated at <2 per cent.

The national Cause-of-Death Register was also scanned. After the sampling point and up to 1985, 1,759 individuals had died. Their underlying causes of death by the ICD-8 classification were retrieved, including definite suicides (E950–959) and deaths where suicide could not be determined with certainty (E980–989). No matching errors occurred.

These data were merged in a new file using the personal identification numbers, which were then removed, to eliminate the identifications. Permission to match these registries was given by the Security Council of Statistics Sweden, the National Board of Health and Welfare, the Data Inspection Board, and the Ethical Review Committee of the Department of Social Welfare, thus meeting with regulatory requirements and suggested ethical practice.

Subgroup characteristics displayed in Tables were analyzed by multiple logistic regression, using the CATMOD procedure in SAS.[17] The dependent variable was either reported regular psychoactive drug use among men and women, respectively, or no or occasional psychoactive drug use among those reporting severe anxiety. Clinically relevant variables were chosen to avoid confounding. To study possible overlap of predictive power and confounding between key variables, the models were repeated without three variables (long-standing illness, disability or infirmity, long-standing nervous symptoms and regular medical treatment). To find "false positive" correlations, models were run also on those reporting good health. Odds ratios and confidence intervals are given for all variables in the models.[18]

Results

The Interview Sample

Regular psychoactive drug use, including prescribed sedatives, neuroleptics, lithium and antidepressants but not hypnotics, was reported by 1.6 per cent of the men and 2.9 per cent of the women (Table 1). The high proportion of elderly among drug users reflects the higher rates of somatic and psychiatric diseases and symptoms, and also concomitant use of analgesics and cardiovascular drugs. The social constrictions may have been related to both age and to poor health: about half the drug users had reduced work capacity, and about one in four had reduced physical mobility.

Only a third of the psychoactive drug users also used hypnotics regularly, although more than half had current sleep problems or fatigue. Nearly all reported some long-standing illness, disability or infirmity, and were receiving

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TABLE 1—Percentage Distribution of Responder Characteristics, Interview Responses, and Registry Data by Sex, Total and Reported Regular Psychoactive Drug Use

Characteristics	Men				Women			
	All interviewed n=12,613 %	Regular drug use n=201 %	Odds ratio	95% CI	All interviewed n=13,708 %	Regular drug use n=394 %	Odds ratio	95% CI
Age >50 years	35	59	1.0*	1.0, 1.0	37	67	1.03*	1.01, 1.06
Not married/cohabitating	34	54	1.0	0.6, 1.5	33	45	0.8	0.6, 1.1
Lives in urban areas	64	70	1.0	0.6, 1.4	65	64	1.0	0.8, 1.4
Basic education only	4	4	1.0	0.4, 2.4	4	3	0.8	0.4, 1.5
Reduced work capacity	10	53	1.6	1.0, 2.6	10	49	1.4	1.0, 2.0
Reduced physical mobility	5	23	0.8	0.5, 1.4	7	30	1.2	0.8, 1.8
No cash margin (SEK 5,000)	14	33	1.2	0.8, 1.9	17	30	1.0	0.7, 1.4
Few or no neighbor relations	49	67	0.7	0.5, 1.1	45	58	0.8	0.6, 1.1
Insomnia in the last two weeks	11	49	1.0	0.6, 1.5	17	60	1.5	1.1, 2.1
Often fatigued in the last two weeks	24	56	1.4	0.9, 2.2	34	61	1.2	0.9, 1.6
Recurrent headaches in last two weeks	8	29	2.8	1.8, 4.5	16	37	1.5	1.1, 2.1
Regular use of hypnotic drugs	1.3	32	5.6	3.1, 10.2	2.6	36	3.9	2.7, 5.8
Regular use of prescribed analgesics	2.6	19	2.5	1.4, 4.2	4.2	23	1.9	1.3, 2.8
Cardiovascular drugs in last two weeks	9	23	1.3	0.6, 2.9	12	32	1.3	0.8, 2.2
Vitamins in the last two weeks	12	25	1.0	0.6, 1.7	21	37	1.4	1.0, 1.8
Long-standing nervous symptoms	2.1	54	20.3	12.4, 33.3	3.2	55	19.3	13.8, 26.9
Long-standing illness, disability	38	97	2.3	0.8, 6.6	41	96	2.4	1.2, 4.7
Regular medical treatment	20	90	9.7	5.0, 18.7	27	88	4.6	2.9, 7.2
Current illness, disability								
Tumor	0.5	1.4	0.9	0.2, 4.2	1.1	2.9	0.7	0.3, 1.7
Endocrine disease	2.6	6	0.5	0.2, 1.1	4.3	14	1.3	0.9, 2.0
Circulatory disease	10	21	0.6	0.4, 1.1	13	31	0.6	0.4, 0.9
Respiratory disease	3.6	7	0.6	0.3, 1.3	3.2	6	0.7	0.4, 1.3
Gastrointestinal disease	4.4	13	0.9	0.5, 1.6	3.8	11	1.0	0.6, 1.6
Musculoskeletal disease	12	23	0.9	0.5, 1.4	12	31	0.9	0.6, 1.2
Other somatic disorders	12	26	1.0	0.6, 1.5	14	28	0.9	0.7, 1.3
Registry data								
Psychiatric admission 1969-83	5.0	44	3.2	1.5, 6.7	4.4	34	2.0	1.2, 3.3
Psychiatric readmission 1969-83	2.8	36	2.9	1.2, 6.7	2.1	23	2.2	1.2, 4.2
Psychiatric substance abuse diagnosis	2.5	17	0.3	0.1, 0.7	0.6	4	0.9	0.3, 2.5
Definite/undetermined suicide	0.3	1.4	1.4	0.2, 10.1	0.1	0.8	0.1	0.0, 0.9

*by age group 16–30, 31–50, 51–74.
Multiple logistic regression models with psychoactive drug use as the dependent variable, by sex. Odds ratios and 95% confidence intervals.

regular medical treatment. Inpatient psychiatric treatment during the 15-year study period was recorded among 44 per cent of the men and 34 per cent of the women who reported regular psychoactive drug use, compared to 5 per cent and 4.4 per cent of all responders. An inpatient psychiatric diagnosis of substance abuse was on record for 17 per cent of the men and 4 per cent of the women, compared to 2.5 per cent and 0.6 per cent of all responders.

In the multiple regression models, the interview and register variables most strongly associated with regular psychoactive drug use were long-standing nervous symptoms, regular medical treatment, regular use of hypnotics and prescribed analgesics and inpatient psychiatric treatment (Table 1).

Anxious Responders

Approximately half of the entire sample was asked a separate question about anxiety in the last two weeks in the 1980 and 1981 interviews. Only 31 per cent of those with severe anxiety reported regular psychoactive drug use; these were primarily women. Those women reporting severe anxiety but no/occasional psychoactive drug use can be characterized as: living alone in urban areas, low income and poorly educated with reduced work capacity, few or no neighborly contacts, and a variety of psychosomatic symptoms (insomnia, fatigue, headaches and nervous symptoms) (Table 2). Three-fourths of the women reported long-standing illness, disability or infirmity (Table 2); most of them suffered from fatigue and insomnia; 21 per cent had been treated in inpatient psychiatric care during the 15-year study period. In another such model of 107 anxious responders who had regular drug treatment, a similar profile was not found.*

Psychiatric Inpatients

The 1,765 psychiatric inpatients were analyzed by their interview responses regarding drug use and symptoms. Only 37 per cent of the drug users had been psychiatric inpatients. Neurotic disorder was the most frequent diagnosis among regular users (Table 3). Of those 137 responders who had received an inpatient psychiatric diagnosis of *schizophrenia*, 64 per cent reported long-standing nervous symptoms, but only 34 per cent reported regular psychoactive drug use. Of those 119 with a diagnosis of affective disorder, only 33 per cent reported regular drug use, yet 42 per cent reported long-standing nervous symptoms.

*Data available on request to author.

TABLE 2—Percentage Distribution of Interview Responses and Registry Data in Those Responders Who Reported Severe Anxiety but no or Occasional Psychoactive Drug Use

Characteristics	n = 243 %	Odds ratio	95% CI
Woman	66	1.6	1.2, 2.2
Age >50 years	48	1.0*	1.0, 1.1
Not married/cohabitating	52	0.7	0.5, 0.9
Lives in urban areas	73	1.5	1.1, 2.1
Basic education only	8	1.9	1.1, 3.2
Reduced work capacity	34	1.8	1.2, 2.8
Reduced physical mobility	18	1.3	0.8, 2.1
No cash margin (SEK 5,000)	33	1.3	0.9, 1.9
Few or no neighbor relations	58	1.4	1.0, 1.8
Insomnia in the last two weeks	61	3.6	2.6, 5.0
Often fatigued in the last two weeks	74	2.7	2.0, 3.8
Recurrent headaches in the last two weeks	39	2.1	1.5, 2.9
Regular use of hypnotic drugs	11	1.7	0.9, 3.0
Regular use of prescribed analgesics	11	0.9	0.5, 1.6
Cardiovascular drugs in the last two weeks	20	0.7	0.3, 1.6
Vitamins in the last two weeks	32	1.3	0.9, 1.8
Long standing nervous symptoms	24	11.0	6.9, 17.4
Long-standing illness, disability	77	1.2	0.7, 1.9
Regular medical treatment	55	1.2	0.8, 1.8
Current illness disability			
Tumor	2.0	1.3	0.5, 3.6
Endocrine disease	7	1.1	0.6, 2.0
Circulatory disease	23	1.0	0.7, 1.6
Respiratory disease	7	1.0	0.5, 1.8
Gastro-intestinal disease	9	0.9	0.5, 1.6
Musculo-skeletal disease	25	1.0	0.7, 1.4
Other somatic disorders	29	1.4	1.0, 2.1
Registry data			
Psychiatric admission 1969-83	21	1.6	0.8, 3.0
Psychiatric readmission 1969-83	13	1.8	0.8, 4.2
Psychiatric substance abuse diagnosis	9	1.7	0.7, 4.0
Definite/undetermined suicide	0	0.1	0.0, 1.3

*by age group 16–30, 31–50, 51–74.

Multiple logistic regression model with no or occasional drug use as the dependent variable: Odds ratios and 95% confidence intervals.

Sixty-three of 323 psychiatric diagnoses (20 per cent) among the psychoactive drug users were for substance abuse disorders, compared to 338 of the 1,025 diagnoses (33 per cent) among all survey respondents. Only 38 individuals had been admitted to inpatient psychiatric care for prescription drug abuse; 19 of them had a concomitant diagnosis of alcohol abuse, and two of street-drug abuse. Thus, only 17 individuals had been admitted for primary prescription drug abuse, a rate of 3.5 per 100,000 person years.

Representativity, Validity and Confounding

Non-responders had received inpatient psychiatric care twice as often as responders (8.1 vs 4.7 per cent) and had higher rates of most psychiatric disorders (Table 3). Since current illness was the reason for non-participation in the interviews in <1 per cent, the difference cannot be explained by illness at the time of the interview. The non-responders also had a higher suicide rate (22/5,727 vs 59/26,952).

The internal validity of the regression models was tested by creating models for those who reported no major health problems (7,337 men and 7,377 women, 55 per cent of all responders). Only two inverted associations to good health (to fatigue and regular hypnotic drug use) among men were found. Thus, the models appear to reflect actual associations between the independent variables and drug use.

TABLE 3—Psychiatric Diagnoses in Inpatient Care per 1,000 Persons among Sample Responders, Non-responders, and Regular Psychoactive Drug Users 1969-83

Diagnoses	All Responders n = 26,952	Non-responders n = 5,727	Regular Drug Users n = 595
Neurosis (n =716)	20.1	30.4	176.5
Alcoholism (n =519)	13.7	26.2	73.9
Unspecified psychosis (n =202)	5.4	9.8	60.5
Schizophrenic psychosis (n =180)	4.3	11.0	65.5
Personality disorder (n =166)	4.6	7.3	45.4
Affective disorder (n =150)	4.4	5.4	65.5
Street-drug abuse (n =100)	2.5	5.8	20.2
Dementia (n =59)	1.3	4.2	15.1
Prescription drug abuse (n =38)	0.9	2.4	11.2
Other psychiatric diagnoses (n =39)	0.8	3.1	8.4

The same person may appear with more than one diagnosis

The external validity of our data appeared satisfactory when compared to the rates of major psychiatric disorders reported in other studies[1,19] and to the disease spectrum among psychoactive drug users.[4-7] The validity of symptoms of fatigue, insomnia, nervousness, and depression reported to lay interviewers was similar to that in another Swedish health survey.[3] In a Finnish health interview survey by nurses, 20 per cent of those reporting both a chronic illness and psychological distress used psychoactive drugs, vs 5 per cent of those who had only psychiatric symptoms.[10] In multiple logistic regression models, the Finnish investigators showed that chronic morbidity, psychiatric symptoms, and physician visits were the best predictors of psychoactive drug use, age being of importance only among women. Our data point to the same conclusion.

When we repeated the regression models without three potentially confounding variables (long-standing nervous symptoms, long-standing illness, disability or infirmity, and regular medical treatment), new associations were noted among drug-using men (age, insomnia, fatigue, being single, having a reduced physical mobility, and lacking neighbor relations), and among drug-using women (fatigue).*

Two of our findings appeared to be spurious. The reduced odds of substance abuse in male drug-takers may be the result of competition with other independent variables. The male substance abusers scored high on a number of indicators of poor health, compared to all males (data to be published). The contribution of those variables (reduced work capacity in 28 per cent of male substance abusers, long-standing nervous symptoms 14 per cent, insomnia 32 per cent, regular psychoactive drug use 11 per cent) to the model reduced the odds ratio of substance abuse, in spite of the fact that 17 per cent of regular psychoactive drug users were also substance abusers. The reduced odds of suicide in female drug-takers may have a similar explanation.

Mortality

In those using psychoactive drugs regularly, there were 79 deaths against 46 expected (95% confidence interval of the difference 15, 51). Among those who reported severe anxiety

there were 23 deaths vs eight expected (95% CI of difference 5, 25). Those diagnosed with schizophrenia had an excess of 19 deaths vs eight expected (95% CI of difference 2, 20), neurotic disorders 34 vs 18 expected (95% CI of difference 4, 28), and with affective disorders there were 11 deaths vs seven expected (95% CI of difference −3, 11).

Among the natural causes of death, only endocrine disorders were more common than expected among the regular psychoactive drug users.

Discussion

Automated record-linkage was used in this study to describe a profile of those in the general population who receive regular treatment with psychoactive drugs. To study causation, longitudinal data are preferred, yet cross-sectional data may contribute to the formulation of explanatory hypotheses for further study.[11]

In addition to the ambiguities inherent in health surveys, the psychoactive drugs could not be subdivided into therapeutic categories. Inpatient psychiatric treatment prior to 1969 as well as any outpatient psychiatric treatment could not be studied. Yet the findings derived with this method appear to be both externally and internally valid. Considering the chronic course of many psychiatric disorders, it is disconcerting that only 42 per cent of those with long-standing nervous symptoms and only one-third of those with schizophrenia were medicating regularly. Another concern is that those who reported severe anxiety, but no or occasional psychoactive drug use were primarily lower-educated women in urban areas, socially isolated and with several symptoms of mental ill health. One wonders whether or not they are in outpatient treatment.

For those who had received psychiatric inpatient treatment, substance abuse was less frequent among drug-users than non-drug-users. One tentative explanation is that pharmacotherapy in patients properly evaluated reduces the risk of substance abuse. The number of prescription drug abusers was small in this sample, indicating that the risk of iatrogenic psychoactive drug abuse leading to inpatient psychiatric care is less common than is generally believed.[21] The assumptions about psychoactive drug overmedication of Swedish general practice patients has also not been shown to hold.[22,23]

Among those admitted for inpatient psychiatric care, the rates of different psychiatric disorders were higher in the non-responders to the health survey, and this was not due to current illness. Although based only on hospital admissions, interpretations of health surveys should consider this source of bias in determining rates of psychiatric morbidity in the general population, and their effects on social drift and suicide. Future investigators will also have to cope with the increasing reluctance among the public on privacy grounds to being scrutinized and recorded in computer files.

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APPENDIX

Swedish Survey of Living Conditions

Statistics Sweden regularly conducts interviews in randomized samples of the Swedish population, similar to the surveys initiated in the United States on the basis of the Health Survey Act 1956. The purpose is to provide the government with a decision-base on socioeconomic conditions, perceived health, and health care utilization. The samples are drawn from the Register of the Total Population, including institutionalized persons, excluding those interviewed in previous samples. Subjects are interviewed by about 200 trained interviewers for one hour at their location; 2–3 per cent are interviewed through a key relative, and 10 per cent over the telephone. Errors in coverage and computational procedures are practically negligible. For the samples in this study, the non-response rate due to refusals was 18.8 per cent in 1975, 19.3 per cent in 1977, 13.7 per cent in 1980, and 13.5 per cent in 1981.

Included in the interview are basic demographic data, housing, leisure facilities, access to public transport, children, general health, medical disorders by the World Health Organization classification system, physical and occupational handicaps, health care utilization, smoking, work conditions, education, unemployment, interest in politics and mass media, relations with neighbors and relatives.

The dependent variable in this study was based on the following questions to the responders: "Have you *in the last two weeks* used any of the following medicines?" A number of categories were listed, such as cough medicines, antibiotics, laxatives, etc. Responses were tabulated as follows:

Tranquilizers (drugs to calm your nerves)	n	% Interviewed
Yes, regularly	595	2.2
Yes, occasionally	750	2.8
No	24,976	92.5
Do not know	36	0.1
No answer	634	2.3
Total	26,991	100

The drugs are not exemplified by brand names, nor is "regular" defined beyond current (last two weeks) and regular use as perceived by the responder.

Personal identification numbers are given to all persons living in Sweden. This unique means of identifying an individual can be used to combine person-based data from different computer files. There are national registries for inpatient care, congenital malformations, cancer, causes of death etc.

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Nurses Working in the Field of AIDS to be Honored; Nominations Invited

October 1989 will be observed as AIDS Awareness and Prevention Month. One of the ways in which the US Public Health Service will participate is to honor nurses in America who have exhibited courage and compassion for people with AIDS and HIV infection in the exercise of their professional duties. Assistant Secretary for Health James O. Mason, MD, DrPH, will present Awards to 10 deserving nurses.

In order to identify these 10 recipients, Dr. Mason has invited nominations from various organizations. The deadline for such nominations is August 15, 1989.

The broad criteria established for these awards is that the individual must be a Registered Nurse with a minimum of three years' experience as a direct care provider for people with AIDS, either as part of his or her usual job or as a volunteer. Although the awards are not limited to US citizens, the work thus recognized must have been performed within the territorial US. Extra consideration will be given for participation in such things as community involvement/outreach programs, recruitment, special care programs, professional organization involvement, publications, and other innovative efforts to ensure care for people with AIDS.

An awards ceremony will be held at the Hubert H. Humphrey Building in Washington, DC on October 4, 1989 at which Dr. Mason will present the awards. PHS will provide transportation for the recipients.

Letters of nomination for the awards, along with a curriculum vitae and a photograph of the nominee, should be sent to: Jim Brown, Director, News Division, Office of Communications, USPHS, 200 Independence Avenue, SW, Room 717H, Washington, DC 20201. Tel: (212) 245-6867.

Alcohol Consumption and Problem Drinking in White Male Veterans and Nonveterans

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Abstract: The relation between veteran status and 15 indicators of alcohol consumption and problem drinking is examined using data from the 1977, 1983, and 1985 National Health Interview Surveys conducted by the National Center for Health Statistics. After simultaneous adjustment for age, region of the US, urbanization, ethnicity, marital status, education, and income, a greater proportion of White male veterans than White male nonveterans were heavy drinkers and a smaller proportion were abstainers. This association appeared in all three survey years; further, veterans consistently reported a high daily consumption level (the equivalent of approximately six bottles of beer, five glasses of wine, or four mixed drinks) compared to nonveterans.

Among White male veterans, 6.9 percent reported family or marital problems and 13.9 percent reported motor vehicle crashes or violations compared to 4.4 percent and 8.9 percent, respectively, among nonveterans. The same pattern was observed for lifetime prevalences of alcoholism (veterans have nearly two and a half times the lifetime prevalence of nonveterans), cirrhosis of the liver, and unspecified alcohol-related problems. (*Am J Public Health* 1989; 79: 1011–1015.)

Introduction

Although there have been numerous studies of alcohol use by men who served in the armed forces in Vietnam, there are few studies of alcohol use among men with other military service experience. A study of men on active duty in stateside Army installations[1] revealed that both officers and enlisted men had larger proportions of heavy drinkers than the civilian population, even after adjustment for age and education. Problem drinking was more common among enlisted men than civilians, an effect which was not apparent among the officers. Cahalan, *et al*, suggested that, ". . . there is something about the Army that fosters drinking over and above those forces at work in the society at large."[1]

In contrast, Polich[2] examined problem drinking in men on active duty in the Army, Navy, and Air Force. Although the problem drinking rate for the military group was somewhat higher than the civilian group, the difference was substantially reduced by adjustment for age, education, and marital status.

A more salient issue is the post-military use of alcohol. Drinking behaviors acquired during military service may persist well into civilian life and can affect an individual's physical or psychosocial well-being. A large body of data suggests that there is a high prevalence of alcohol problems and alcoholism among Veterans Administration patients.[3–6] These "rate-under-treatment studies" are problematic, however, because of the unique characteristics of help-seeking veterans. Boscarino[7,8] surveyed alcohol consumption in a random sample of veterans and nonveterans and concluded that the higher prevalence of heavy drinkers among veterans is accounted for by differences in nine demographic characteristics: sex, age, race, region of United States, urbanization, ethnicity, marital status, education, and income.

The variable conclusions of the Cahalan, *et al*, Polich, and Boscarino surveys may be in part due to differences in the populations surveyed (men on active duty vs veterans), the outcomes measured (excessive consumption vs problem drinking vs some construct of the two), and the extent and type of demographic adjustment performed (single variable adjustment vs multivariable adjustment). The purpose of this investigation is to examine alcohol consumption and problem drinking in White male veterans and nonveterans who participated in the National Health Interview Survey Supplements of 1977, 1983, and 1985. To test the hypothesis that White male veterans drink more than White male nonveterans and experience more alcohol-related adverse consequences, we examined 15 alcohol use outcomes after adjustment for demographic characteristics.

Methods

Sources of Data

The National Health Interview Survey (NHIS) is an annual household survey by the National Center for Health Statistics (NCHS) involving personal interviews of a nationwide sample of the civilian, noninstitutionalized population of the United States. The sampling plan of the survey follows a multistage probability design which ensures that these households are representative of the target population. Details of the NHIS are presented fully elsewhere.[9]

On an irregular basis, the NCHS adds special supplemental components to the core NHIS questionnaire. The supplemental questionnaires administered in 1977, 1983, and 1985 elicited information on alcohol use. Unlike the core questionnaire which relies on cost-effective proxy responses, all three of these supplements use direct self-reporting to obtain more accurate information.

Measures of Veteran Status

Veteran status is ascertained in the core NHIS questionnaire in 1977, 1983, and 1985. Using computer tapes for the three interview years, we identified male veterans and male nonveterans to obtain the following sample sizes: in 1977, 4,005 veterans and 5,845 nonveterans; in 1983, 2,901 veterans and 6,004 nonveterans; and in 1985, 4,079 veterans and 8,998 nonveterans. However, we limited our selection to those veterans and nonveterans with complete data for the outcome variable and the seven covariates (age, region of the United States, urbanization, ethnicity, marital status, education, and income) used in the multivariable adjustment procedure. Our definition of complete data led to the exclusion of such survey responses as "unknown," "refused," or "other." This selection reduced the 1977 sample by 28

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percent, the 1983 sample by 13 percent and the 1985 sample by 11 percent. We further restricted the sample to White male veterans and nonveterans, recognizing that persons of other races tend to have different drinking patterns than Whites but contribute little in the way of sample size. Thus, the final sample sizes for the three interview years are: 1977, 2,707 veterans and 3,704 nonveterans; 1983, 2,326 veterans and 4,590 nonveterans; and 1985, 3,181 veterans and 6,749 nonveterans.

Measures of Alcohol Consumption

From the alcohol data included on the NHIS supplements we selected 15 dependent variables which provide a concise description of drinking behavior. The 15 variables can be grouped into two categories representative of alcohol use: level of consumption, and adverse consequences. The variables which delineate adverse consequences of alcohol use include five adverse social consequences and seven adverse health consequences (see Appendix A). Owing to the varying amount of detail in each supplement year with regard to alcohol use questions, these 15 outcomes are associated with data from one, two, or three supplement years as shown in Appendix A. The differences between these measures according to survey year are discussed in Appendix B.

Statistical Analysis

The statistical analysis of the relationship between veteran status and the 15 alcohol use outcomes employs a logistic regression procedure to produce predicted proportions of veterans and nonveterans in each category of the dependent variable. If the outcome is ordinal, a modification of the logistic model for dichotomous outcomes is used.[10] Although it is possible to conduct these analyses using the sampling weights provided by NCHS, we chose to perform an unweighted analysis owing to the fact that our comparison category (veteran vs nonveteran) is not influenced by sampling.

Logistic models for each of the 15 outcomes are constructed to examine the effect of veteran status. These models include seven potential confounders: age, region of the US, urbanization, ethnicity, marital status, education, and income. By examining the adjusted proportions we can determine what drinking behaviors, if any, are associated with prior military service above and beyond the more general societal influences. The adjusted proportions are produced using a covariance adjustment method.[11,12]

Results

Levels of Alcohol Consumption

The top portion of Table 1 presents the relationship between veteran status and drinking level after adjusting for age, region of the US, urbanization, ethnicity, marital status, education, and income. In each supplement year, a positive association is observed between drinking level and veteran status. Among nonveterans, there are more abstainers and light drinkers whereas among veterans, there are more moderate and heavy drinkers. The proportion of heavy drinkers among veterans consistently exceeds the proportion of heavy drinkers among nonveterans by 2 to 6 percent.

As with drinking level, a positive association between high alcohol consumption and veteran status is evident in all three survey years (Table 1). This association is most pronounced in 1983, where the proportion of veteran high consumers is nearly twice that of nonveteran high consumers. If the 1983 high consumer were a single beverage drinker, that individual would report a daily intake equivalent to approximately six bottles of beer, five glasses of wine, or four mixed drinks.

In all survey years, a larger proportion of nonveterans than veterans report that they have not had five or more drinks in one day during the past 12 months. Conversely,

TABLE 1—Levels of Alcohol Consumption in White Male Veterans and Nonveterans: Drinking Level, High Consumption and Number of Heavy Consumption Days[a]

	Interview Year and Veteran Status					
	1977		1983		1985	
	% Veterans	% Nonveterans	% Veterans	% Nonveterans	% Veterans	% Nonveterans
Drinking Level[b]	(n = 2,707)	(n = 3,704)	(n = 2,326)	(n = 4,590)	(n = 3,181)	(n = 6,749)
Abstainer	10.5	12.9	15.8	21.0	23.8	28.4
Light	24.7	27.7	25.8	29.3	18.2	19.5
Moderate	44.1	42.2	34.6	31.7	38.5	36.1
Heavy	20.7	17.2	23.8	18.0	19.5	16.0
High Consumption[c]	(n = 2,209)	(n = 2,906)	(n = 1,399)	(n = 2,750)	(n = 1,969)	(n = 4,430)
No	87.9	90.6	90.2	94.4	89.2	93.5
Yes	12.1	9.4	9.8	5.6	10.8	6.5
Number of Heavy Consumption Days[d]			(n = 1,697)	(n = 3,274)	(n = 1,914)	(n = 4,271)
0 days			38.5	42.1	27.8	36.2
1–3 days			15.7	15.8	12.7	13.9
4–11 days			16.3	15.6	18.1	17.5
≥ 12 days			29.5	26.5	41.4	32.4

[a]Percentages are adjusted for age, region of the US, urbanization, ethnicity, marital status, education, and income. These variables are described in detail in Appendix A.
[b]Measured on the entire population of veterans and nonveterans.
[c]Measured on current drinkers only.
[d]Measured on current and noncurrent drinkers. Note that this information was not obtained in the 1977 interview.

TABLE 2—Adverse Consequences of Alcohol Consumption in White Male Veterans and Nonveterans: 1983 NHIS Lifetime Social Consequences[a]

Social Consequences	Veteran Status	
	% Veterans	% Nonveterans
Family or Marital Problems	(n = 2,064)	(n = 3,739)
no problems	93.1	95.6
1 or more problems	6.9	4.4
Motor Vehicle Crash or Violation	(n = 2,059)	(n = 3,734)
no incidents	86.1	91.1
1 or more incidents	13.9	8.9
Job Problems	(n = 2,063)	(n = 3,739)
no problems	97.9	98.2
1 or more problems	2.1	1.8
Injury	(n = 2,063)	(n = 3,737)
no injuries	95.6	95.9
1 or more injuries	4.4	4.1
Social Consequences Score	(n = 2,059)	(n = 3,729)
0 problems	79.6	85.7
1–2 problems	14.2	10.1
≥ 3 problems	6.2	4.2

[a]Percentages are adjusted for age, region of the US, urbanization, ethnicity, marital status, education, and income. These variables are measured on current, noncurrent, and former drinkers, and are described in detail in Appendix A.

veterans are more likely than nonveterans to report 12 or more heavy consumption days in the past year (Table 1).

Adverse Consequences of Alcohol Consumption

The relation of veteran status and lifetime adverse social consequences of alcohol consumption adjusted for demographic factors is presented in Table 2. Among veterans, the proportion of men experiencing family and marital problems is one and a half times the size of the comparable proportion among nonveterans; a similar pattern is observed for alcohol-related motor vehicle crashes and violations. Veteran status appears to be unrelated to injury. The additive social consequences score is elevated for veterans.

TABLE 3—Adverse Consequences of Alcohol Consumption in White Male Veterans and Nonveterans: 1983 NHIS Lifetime Health Consequences[a]

Health Consequences	Veteran Status	
	% Veterans	% Nonveterans
Alcoholism	(n = 2,067)	(n = 3,741)
no	93.6	97.4
yes	6.4	2.6
Cirrhosis of the Liver	(n = 2,065)	(n = 3,760)
no	99.3	99.9
yes	0.7	0.1
Unspecified Alcohol-related Health Problem	(n = 2,055)	(n = 3,734)
no	95.2	96.8
yes	4.8	3.2
Delirium Tremens	(n = 2,070)	(n = 3,749)
no	98.6	98.9
yes	1.4	1.1
Blackouts	(n = 2,070)	(n = 3,762)
no	93.9	93.2
yes	6.1	6.8
Fatty Liver	(n = 2,068)	(n = 3,758)
no	99.3	99.6
yes	0.7	0.4
Health Consequences Score	(n = 2,042)	(n = 3,703)
0 problems	88.2	90.2
1–2 problems	9.1	7.6
≥ 3 problems	2.7	2.2

[a]These variables are measured on current, noncurrent and former drinkers, and are described in detail in Appendix A. Percentages are adjusted for age, region of the US, urbanization, ethnicity, marital status, education, and income.

Table 3 presents the relationship between veteran status and lifetime adverse health consequences of alcohol consumption after adjusting for demographic factors. These results are not as consistent as those for the social consequences, involving fewer events and less stable estimates. Event categories involving less than five cases are not reported. Compared to nonveterans, veterans report an increased lifetime prevalence of alcoholism, cirrhosis of the liver, unspecified alcohol-related health problems, and all health consequences combined. Of these four categories, the alcoholism outcome is the most compelling: veterans have nearly two and a half times the lifetime prevalence of alcoholism compared to nonveterans.

Discussion

A fundamental concern in cross-sectional studies of alcohol use is the issue of adequate control for confounding factors. Certainly our adjustment for seven demographic variables cannot be termed exhaustive for it does not include any pre-service (e.g., parents' drinking habits) or in-service (e.g., combat exposure) variables.

An informative comparison can be made between our work and the study by Boscarino.[7] Boscarino examined level of alcohol consumption in a 1975 national probability sample of 3,071 American adults. As with the NHIS survey, veteran status was ascertained by self-report.

Although Boscarino does not provide specific details as to how his ethanol consumption estimate was derived, he does state that it is based on the "quantity and frequency of 'usual' alcoholic beverages consumed in the 30 days preceding the interview." Our 1983 estimate relies on exact quantities and frequencies of each beverage type in the 14 days prior to the interview. Nonetheless, in his unadjusted results for level of consumption, Boscarino found that veterans are significantly more likely than nonveterans to report higher consumption levels.

In the examination of adjusted results, however, our conclusions differ from those of Boscarino. When Boscarino adjusted for sex, race, age, region of the United States, urbanization, ethnicity, marital status, education, and income the excess percent of heavier drinkers among veterans disappeared. Boscarino replicated his 1980 findings using a second survey conducted in 1977 (n = 3,322 adults).[8]

What might explain the discrepancy between Boscarino's and our results? It is unlikely that the divergent results are due to different alcohol consumption measures since both studies use similar measures. However, our analytical procedures for adjustment differ from Boscarino's in several respects. First, Boscarino includes both males and females in his sample. He attempted to adjust for potential confounding by sex using a regression model which included sex as an independent variable. This is a questionable strategy because women (who have very different alcohol consumption levels from men) comprise only 2 percent of Boscarino's veteran sample but approximately 50 percent of the nonveteran sample. We excluded females from our analysis because few females are veterans. Second, Boscarino's data are composed of both Whites and non-Whites and race is included as a dichotomous adjustment factor. Recognizing that the "non-White" group (which includes Blacks, Asians, and other categories) is heterogeneous with respect to alcohol use, we restricted our analysis to Whites. Third, Boscarino's use of

ordinary least squares regression with an ordinal dependent variable is not entirely appropriate. We used a form of logistic regression adapted for ordinal dependent variables. Fourth, one of Boscarino's nominal independent variables (urbanization) was not dummy coded but was left in its original form. The presumption of a linear increase in risk of consumption implied by such coding is questionable. We dichotomized all of the independent variables except for age, education, and income, which were coded ordinally.

The 1988 NHIS supplement includes questions on alcohol use as well as the standard veteran question. When these data become available, it will be possible to determine if some of the associations demonstrated here between veteran status and alcohol use are replicated.

APPENDIX A
Description of Alcohol Use Variables Analyzed

Category	Dependent Variable	Supplement Years Available	Subpopulation Referenced[a]
Level of Consumption	Drinking level[b]	1977, 1983, 1985	Total
	High consumption[c]	1977, 1983, 1985	Current drinkers
	Heavy consumption days[d]	1983, 1985	Current and noncurrent drinkers
Lifetime Adverse Consequences of Alcohol Use	*Social Consequences*	1983	Current, noncurrent, and former drinkers
	Family or marital problems		
	Motor Vehicle crash or violation		
	Job problems		
	Injury		
	Social consequences score[e]		
	Health Consequences	1983	Current, noncurrent, and former drinkers
	Alcoholism		
	Cirrhosis of the liver		
	Unspecified alcohol-related health problem		
	Delirium tremens		
	Blackouts		
	Fatty liver		
	Health consequences score[f]		

a) The total population includes current drinkers, noncurrent drinkers, former drinkers, and abstainers. Current drinkers report having had a drink in the two weeks preceding the interview whereas noncurrent drinkers had their last drink more than two weeks (but less than a year) prior to the interview. A former drinker's last drink was more than a year before the interview; an abstainer has had less than 12 drinks in any one year of his life.

b) Drinking level is based on a composite measure which reflects the average daily ethanol consumption (ADEC) from all alcoholic beverages (beer, wine, and liquor) combined. ADEC is calculated from the Quantity/Frequency (Q/F) items of the questionnaires and then categorized to obtain drinking level. The levels correspond to NIAAA guidelines where: abstainers average 0.0 ounces ethanol daily; light drinkers, 0.01–0.21 ounces ethanol daily; moderate drinkers, 0.22–0.99 ounces ethanol daily; and heavy drinkers, 1.00 or more ounces of ethanol daily.

c) High Consumption is also based on the composite measure ADEC. High consumption is simply the number of ounces of ethanol corresponding to the upper fifth percentile of the ADEC distribution for veteran and nonveteran drinkers combined. These "cutoff" levels of consumption are specific for interview year and represent the observed extreme among current drinkers. In 1977, this extreme is 1.96 ounces of ethanol per day. In 1983, it is 2.88 ounces of ethanol per day, and in 1985, 2.14 ounces of ethanol per day.

By converting ounces of ethanol to ounces of beer, wine, or liquor, it is possible to determine how many drinks constitute high consumption.[13] For example, high consumption in 1983 corresponds to six or more beers per day:

$$\frac{1.00 \text{ oz. beer}}{0.04 \text{ oz. ethanol}} \times \frac{1 \text{ bottle}}{12 \text{ oz. beer}} \times 2.88 \text{ oz. ethanol/day} = 6.0 \text{ bottles/day}$$

Similar equations can be set up for wine and liquor in the 1983 data set:

$$\frac{1.00 \text{ oz. wine}}{0.15 \text{ oz. ethanol}} \times \frac{1 \text{ glass}}{4 \text{ oz. wine}} \times 2.88 \text{ oz. ethanol/day} = 4.8 \text{ glasses/day}$$

$$\frac{1.00 \text{ oz. liquor}}{0.45 \text{ oz. ethanol}} \times \frac{1 \text{ drink}}{1.5 \text{ oz. liquor}} \times 2.88 \text{ oz. ethanol/day} = 4.3 \text{ drinks/day}$$

These calculations can be repeated for the 1977 and 1985 data sets.

d) The number of heavy consumption days is the number of days in the past year when the respondent reports having consumed five or more alcoholic drinks in one day.

e) This score is a simple summation of the number of social problems (family/marital, motor vehicle crash/violation, job, injury) reported by the respondent. Respondents indicate whether or not they have ever experienced a drinking-related event involving their family or spouse, a motor vehicle crash or violation, their job, or an injury. (For example, the injury question reads, "Have you EVER had an injury related to YOUR drinking?".[5]) An overall score was not computed for a respondent if any problem item was missing; thus, the sample size for the social consequences score is smaller than the sample sizes associated with the individual problems.

f) This score is a simple summation of the number of health problems (alcoholism, cirrhosis of the liver, unspecified alcohol-related health problem, delirium tremens, blackouts, fatty liver) reported by the respondent. Respondents are asked about the cumulative health effects of alcohol including alcoholism, cirrhosis of the liver, an unspecified alcohol-related health problem, delirium tremens, blackouts, or fatty liver. Unlike the social consequences, the health consequences are not specifically related by the interviewer to alcohol use; rather, respondents are simply asked if they have ever had the condition. We chose from a list only those conditions which are the prima facie endpoints of chronic, excessive alcohol use. As with the health consequences score, the sample size for the social consequences score is smaller than the sample sizes associated with the individual problems.

APPENDIX B
Details of the Differences in the 1977, 1983, and 1985 NHIS Supplemental Interviews

Definitions of Drinkers

The 1983 questionnaire includes groups of questions written especially for subjects who currently drink, subjects who are not currently drinking, former drinkers, and abstainers. Current drinkers report having had a drink in the two weeks preceding the interview whereas noncurrent drinkers had their last drink more than two weeks (but less than a year) prior to the interview. A former drinker's last drink was more than a year before the interview; an abstainer has had less than 12 drinks in any one year of his life.

In the 1977 and 1985 questionnaires, however, the majority of the alcohol questions are asked of current (and occasionally noncurrent) drinkers only. The subpopulation referenced for each variable in Appendix A represents not only that variable's availability for a specific type of drinker but an attempt on our part to maximize the interpretability and consistency of that measure across supplement year. In general, the level of consumption variables measure "current" drinking or, in the case of heavy consumption days, drinking over the past year. The adverse consequence variables, on the other hand, refer to drinking problems experienced over a lifetime.

Definition of Alcohol Use Measures

Drinking level and ADEC are generated by NCHS in the 1983 data but must be calculated for the 1977[14] and 1985[15] data sets. Moreover, it should be noted that the 1977 and 1985 measures of drinking level are only approximations of the 1983 measure. For example, in 1977 respondents were asked if they drank beer (or wine or liquor) "never," "occasionally," "once or twice a week," "more than twice a week," or with some other frequency. This is in contrast to the 1983 frequency question which requires the respondent to count the number of days out of the last 14 in which beer (or wine or liquor) was consumed. The 1985 Quantity/Frequency questions ask not about specific beverages but about "drinks" in general; an assumption about the ethanol content of an unspecified drink must then be made.[15] Thus, an examination of drinking level or high consumption across supplement year must be done with caution. The heavy consumption days question (asked only of 1983 and 1985 respondents) is consistent between questionnaires, however.

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International Conference: Public Health and Sexual Transmission of Diseases

An international meeting—Public Health and the Sexual Transmission of Diseases: Directions for Future Research and Health Policy—is being organized to take place in Chamonix, France, April 23–25, 1990. Limited to only 140 participants, the proposed scientific meeting will focus on those STDs of major public health importance: infections due to gonococcus and chlamydia trachomatis, papilloma virus, and human immunodeficiency virus (HIV). The aim of the meeting is to study these diseases from a public health perspective, adopting a common approach to the problem of AIDS (acquired immunodeficiency syndrome) or other STDs. Focusing on the relation between behavior and the development of STDs, including AIDS, and on preventive measures to reduce the incidence and restrict the consequences of these diseases, the meeting will include four plenary sessions:

- STDs: A challenge to public health
- STDs: A challenge to society
- Individual and group behavior and the spread of STD
- STD prevention: Strategies and evaluation

A free paper session and two round tables, "Directions for future research" and "Recommendations for future public health initiatives," will complete the program.

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A Case-Control Study of Pancreatic Cancer and Cigarettes, Alcohol, Coffee and Diet

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Abstract: A pancreatic cancer case-control study was conducted in the Minneapolis-St. Paul area. Family members were interviewed about the subject's usage of cigarettes, alcohol, coffee, and other dietary factors in the two years prior to death (cases, n = 212) or prior to interview (controls, n = 220). The adjusted odds ratio for two packs or more of cigarettes per day was 3.92 (95% CI = 1.18, 13.01) and four or more drinks per day OR 2.69 (95% CI = 1.00, 7.27). Coffee was not a risk factor (seven cups or more per day; OR 0.58 (95% CI = 0.27, 1.27). A positive trend was observed for beef and pork consumption, and a negative trend from cruciferous vegetables. (*Am J Public Health* 1989; 79:1016–1019.)

Introduction

Cigarette smoking is the only currently known etiologic factor for pancreatic cancer.[1] While some epidemiologic studies have shown a positive association between alcohol consumption and pancreatic cancer,[2–4] most have not.[5–11] A positive association between pancreatic cancer and coffee consumption was reported in one study,[5] but was not substantiated by other studies.[6–10] Four case-control studies have reported inconsistent findings regarding meat and vegetable consumption.[7–10] The present case-control study was designed to evaluate further the role of cigarettes, alcohol, and coffee as well as to explore other dietary factors.

Methods

Cases were White males ages 40 to 84 whose deaths occurred between January 1, 1980 and June 30, 1983 with pancreatic cancer (exocrine only) listed on the death certificate; place of residence at the time of death was within the seven-county Minneapolis-St. Paul area.

There were 262 pancreas-identified cancer deaths. An informant was located for 250 cases and 212 (84.8 percent) were interviewed. Of those interviewed, 140 cases (66.0 percent) had pathological confirmation.

Population-based controls were White males, ages 40–84, ascertained through random digit dialing and frequency matched within five year groups. A total of 94.2 percent of all households screened provided data.[12] The following are the percentages (in parentheses) of eligible controls whose respondents agreed to participate in the study: less than 60 years of age (80.2), ages 60–64 (82.2), 65–69 (92.6), 70–74 (72.7), 75 and older (69.5).

The widow (case) or spouse (control) was interviewed at home. If unavailable, the next closest relative (e.g., daughter) was sought. The spouse was the primary respondent for 73 percent of the cases and 92 percent of the controls. The subject's wife or daughter provided the information for 88 percent of the cases and 97 percent of the controls.

Questions on smoking, drinking, and diet pertained to the subject's usage two years prior to death for the cases and two years prior to the interview for the controls. Individual dietary information was obtained by a frequency-of-use questionnaire. Frequency categories ranged from no use to daily use (two or more packs of cigarettes, four or more drinks of beer, wine, or hard liquor, and seven or more cups of coffee). An index of alcohol consumption was computed using the combined frequencies of beer, wine, and hard liquor. In-season and out-of-season consumption was asked for selected vegetables and fruit. Reported are the the monthly consumption indices for beef, pork, poultry, cruciferous vegetables (cabbage, broccoli, brussel sprouts, and cauliflower) and noncruciferous vegetables, fruits and juices, bread and cereals, and dairy products.

Logistic regression was employed to obtain maximum likelihood point estimates and 95 percent confidence intervals of the odds ratio.[13] Odds ratios were adjusted for the potential confounding variables of age, educational status, reported history of diabetes mellitus, cigarettes, alcohol as well as other dietary life-style factors where applicable, i.e., frequency of meat (beef, pork and poultry) and vegetable (cruciferous and noncruciferous) consumption.

Analyses were performed using all cases and only histologically confirmed cases. Since there were no differences, the results are reported for all cases combined. Data were also analyzed and reported separately for all respondents and spouse only respondents.

Results

An increased risk for pancreatic cancer was observed for current cigarette smoking and, to a less extent, smoking duration (Table 1). An index of total alcohol consumption (Table 2) showed a greater than two-fold risk for pancreatic cancer for heavy alcohol consumption (four or more drinks per day), primarily due to heavy beer and hard liquor consumption. There was weakly decreased risk for coffee consumption. A positive trend was observed for the consumption of beef and pork but not poultry (Table 3). Consumption of cruciferous vegetables showed a negative trend but this was not observed with the consumption of noncruciferous vegetables or fruits and juices. Increased risks were associated with dairy products and breads and cereals. However, a strong difference was observed in the consumption of white versus whole wheat bread. The upper quartile

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TABLE 1—Adjusted Odds Ratios* and 95 Percent Confidence Intervals for Cigarette Smoking by Type of Respondent

Current Smoking Status	All Respondents			Spouses Only		
	Cases/ Controls	Adjusted Odds Ratio	95% CI	Cases/ Controls	Adjusted Odds Ratio	95% CI
Nonsmokers	132/178	1.00	—	94/164	1.00	—
Current Packs/Day						
<1 pack	22/17	1.77	0.87–3.61	16/16	1.68	0.77–3.69
1–2 packs	42/21	2.90	1.55–5.43	33/21	2.83	1.47–5.46
≥2 packs	16/4	3.92	1.18–13.01	11/4	2.65	0.72–6.20
Years smoked**						
1–15	11/6	2.06	0.70–6.05	7/6	1.75	0.54–5.68
16–29	30/16	2.58	1.26–5.28	25/16	2.55	1.21–5.39
≥30	34/19	2.37	1.25–4.49	24/18	2.15	1.06–4.35
Never Smoked	44/54	1.00	—	30/49	1.00	—
Ex smokers	88/124	0.76	0.46–1.28	64/115	0.76	0.43–1.37

*Adjusted for age, education level, reported diabetes mellitus history, and alcohol, meat and vegetable consumption.
**Five spouse and one non-spouse respondents were unable to recall smoking duration.

TABLE 2—Adjusted Odds Ratios* and 95 Percent Confidence Intervals for Total Alcohol, Beer, Hard Liquor, Wine and Coffee Consumption by Type of Respondent

Drinks/Day	All Respondents			Spouses Only		
	Cases/ Controls	Adjusted Odds Ratio	95% CI	Cases/ Controls	Adjusted Odds Ratio	95% CI
Total alcohol						
0	56/44	1.00	—	40/39	1.00	—
1	104/149	0.77	0.47–1.30	77/140	0.73	0.41–1.28
2–3	28/20	1.42	0.67–3.03	20/19	1.30	0.57–2.97
≥4	24/7	2.69	1.00–7.27	17/7	2.33	0.81–6.67
Beer						
0	89/80	1.00	—	63/72	1.00	—
1	100/132	0.87	0.56–1.35	77/125	0.84	0.52–1.36
2–3	10/7	1.09	0.35–3.39	5/7	0.62	0.16–2.31
≥4	13/1	8.25	1.01–67.54	9/1	6.92	0.80–59.52
Hard Liquor						
0	95/78	1.00	—	65/67	1.00	—
1	86/126	0.66	0.42–1.01	64/123	0.60	0.37–0.98
2–3	24/14	1.24	0.56–2.71	19/13	1.34	0.57–3.11
≥4	7/2	3.49	0.65–18.72	6/2	3.53	0.63–19.70
Wine						
0	119/78	1.00	—	85/69	1.00	—
1	89/138	0.63	0.41–0.97	65/132	0.55	0.36–0.88
2–3	2/3	0.69	0.09–5.43	2/3	0.83	0.11–6.40
≥4	2/1	2.72	0.24–31.37	2/1	2.97	0.26–34.13
Coffee						
<1	29/25	1.00	—	19/24	1.00	—
1–3	60/84	0.50	0.26–1.00	42/77	0.59	0.28–1.26
4–6	74/68	0.72	0.37–1.45	54/62	0.83	0.39–1.77
≥7	49/43	0.58	0.27–1.27	39/42	0.68	0.30–1.57

*Adjusted for age, education level, reported diabetes mellitus history, cigarette smoking, meat and vegetable consumption.

of cases ate more white bread (OR 2.42, 95% CI = 1.52, 3.84) and less whole wheat bread (OR 0.44, 95% CI = 0.26, 0.76).

Discussion

The reporting of pancreatic cancer on the death certificate has been shown to be quite accurate.[14] The use of death certificates allowed for the virtually complete ascertainment of all cases in the defined geographic area and justified the use of a population-based control group.

In this study, live rather than deceased controls were selected. The use of deceased controls may introduce a bias since their use of alcohol and cigarettes may be greater than the general population.[15] To keep data collection procedures comparable in cases and controls, we asked surrogates for the information for the controls. Disparaties in recall could still exist, but several features of our method minimized the influence that a live control could have had on a surrogate: the surrogate was not informed about the nature of the interview until the interviewer was present; controls were not allowed to be present during the interview; potentially sensitive questions (e.g., alcohol consumption) were asked toward the end of the interview.

Diseases such as pancreatic cancer which are rapidly fatal are difficult to study because many patients die before the interview can be completed or are too debilitated to participate. Previous case-control studies have suffered from this problem. Wives and daughters can provide reasonable estimates of a man's dietary habits, although the data on alcohol are less certain.[16–18]

The risk estimates and trends observed in this study are consistent with other results for cigarette smoking[2–11] and

TABLE 3—Adjusted Odds Ratio and 95 Percent Confidence Intervals for Selected Food Groups by Type of Respondent

Food Group Times/Month	All Respondents			Spouses Only		
	Cases/Controls	Adjusted Odds Ratio	95% CI	Cases/Controls	Adjusted Odds Ratio	95% CI
Beef						
≤8	43/53	1.00	—	22/50	1.00	—
9–17	89/112	1.05	0.62–1.78	67/106	1.57	0.84–2.92
≥18	80/55	1.81	1.00–3.28	65/49	3.02	1.50–4.89
Pork						
≤2	29/43	1.00	—	21/39	1.00	—
3–8	106/119	1.42	0.78–2.57	80/112	1.43	0.74–2.76
≥9	77/58	1.90	1.00–3.61	53/54	1.66	0.81–3.39
Poultry						
≤2	45/42	1.00	—	28/39	1.00	—
3–5	106/120	0.80	0.47–2.23	83/111	1.09	0.60–1.99
≥6	61/58	0.95	0.52–1.73	43/55	1.09	0.55–2.14
Cruciferous vegetables						
≤2	62/49	1.00	—	38/39	1.00	—
3–8	102/114	0.71	0.43–1.18	72/111	0.56	0.31–1.01
≥9	48/57	0.57	0.31–1.04	44/55	0.55	0.28–1.09
Noncruciferous vegetables						
≤16	71/56	1.00	—	44/51	1.00	—
17–31	78/113	0.55	0.33–0.91	57/104	0.58	0.33–1.03
≥32	63/51	0.95	0.52–1.73	53/50	1.07	0.55–2.06
Fruits and juices						
≤21	68/53	1.00	—	47/49	1.00	—
22–52	98/112	0.90	0.54–1.49	67/102	0.91	0.51–1.61
≥53	46/55	0.88	0.48–1.62	40/54	1.06	0.54–2.06
Breads and cereals						
≤24	32/52	1.00	—	23/48	1.00	—
25–36	95/110	1.22	0.68–2.12	66/104	1.34	0.70–2.58
≥37	85/58	2.19	1.18–4.08	65/53	2.66	1.33–5.33
Dairy						
≤16	50/56	1.00	—	31/50	1.00	—
17–38	97/102	1.21	0.72–2.03	70/94	1.26	0.70–2.27
≥39	65/62	1.51	0.85–2.71	53/61	1.59	0.84–3.01

*Adjusted for age, education level, reported diabetes mellitus history, cigarette smoking, alcohol and when appropriate meat and/or vegetable consumption. The frequency level was determined by the 25th and 75th percentile values of the controls in order to categorize low, moderate and high consumption patterns.

Beef = hamburger, pot roast, roast beef, steak, beef ribs.
Pork = pork chops, ham, ribs, roast, bacon.
Poultry = chicken, turkey.
Cruciferous vegetables = cabbage, broccoli, brussel sprouts, cauliflower.
Noncruciferous vegetables = spinach, carrots, lima beans, peas, tomatoes, corn, green and yellow beans.
Fruits and juices = oranges, apples, bananas, peaches, grapefruit, strawberries, orange juice.
Breads and cereals = white, whole wheat, rye bread, cereal.
Dairy = cheese, cottage cheese, ice cream, milk.

further support its role as an etiologic factor for pancreatic cancer.

Heavy alcohol consumption adjusted for several potential confounding variables was positively associated with pancreatic cancer. The consumption of beer and hard liquor contributed to this association, but there were few wine drinkers among the subjects. These results are supported by some[2–4] but not all epidemiologic studies on pancreatic cancer. This discrepancy, reviewed by Velema, *et al*,[19] may be due, in part, to differences in case selection, hospital versus population-based controls, the use of living or deceased controls when cases are deceased, and whether the vital status of the case affects the accuracy of information provided by a family member. Studies of alcohol abusers have not reported an excess risk for pancreatic cancer;[20,21] however, these studies did not allow for an adequate latency period and many abusers died of other causes.

Heavy alcohol consumption may induce chronic calcifying pancreatitis which may, perhaps in the presence of other dietary modifiers (e.g., high fat), lead to pancreatic cancer.[22] Animal models, however, have not shown pancreatic cancer to be produced by high concentrations of ethanol.[23] At this time there is no clear biologic mechanism to link alcohol consumption to pancreatic cancer.

It is probably premature to suggest, as some investigators have,[1,9] that alcohol does not have an important role in the development of pancreatic cancer. We agree with Velema, *et al*,[19] that the consumption of alcohol in relation to pancreatic cancer should remain under close observation.

Our findings concur with most epidemiologic[6–11] and laboratory[24] studies that coffee is not a risk factor for pancreatic cancer. In fact, after adjusting for major dietary life-style factors as recommended by Falk, *et al*,[10] we observed a negative association for coffee usage even at seven or more cups per day.

Ecologic studies have examined national food consumption patterns and pancreatic cancer mortality rates and have shown a positive correlation between pancreatic cancer and meat and fat consumption;[25,26] of four recent case-control studies, three have reported positive associations with meat. Mack, *et al*, observed a high level of beef consumption among directly interviewed subjects.[9] Falk, *et al*, showed a dose response for pork products.[10] Norell, *et al*, reported a positive association with meat that was fried or grilled.[8] However, Gold, *et al*, in their case-control study, did not find a relation with beef, deep-fried foods, smoked or barbequed

beef, bologna, or salami.[7] Laboratory data have suggested high fat diets may cause pancreatic hyperplasia which may subsequently lead to neoplasia.[27] Our findings support the hypothesis that diets high in animal fat may be associated with pancreatic cancer.

Our results showed a decreased risk of pancreatic cancer with an increased consumption of cruciferous vegetables. Indoles and aromatic isothiocyanates, found in cruciferous vegetables, have been shown to increase glutathione S-transference activity and inhibit experimentally induced neoplasia in laboratory animals.[28] The lack of a decreasing risk trend for consumption of noncruciferous vegetables was due to cases consuming more of the common vegetables (peas, corn, green bean and yellow beans). Unlike Norell, *et al*,[8] we did not observe a negative trend with carrots.

The risk estimates for fruit consumption showed at best a weak negative association. Others have reported much stronger negative associations.[7–10]

The increased risk for white bread has also been observed elsewhere.[7,9] This association may be due to a change from white to whole wheat bread by controls which may parallel their preference for a diet of more vegetables, less meat.

Based on the results of our study and those from previously published studies, we believe a cessation of cigarette smoking will result in a decreased incidence of pancreatic cancer in the adult male population. Furthermore, a decrease in heavy alcohol and meat consumption and a corresponding increase in vegetables may also contribute to a decline. Our results lend additional support to the conclusion that coffee is not a risk factor for pancreatic cancer.

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Call for Abstracts

APHA Late-Breaker Epidemiology Exchange Session

The Epidemiology Section will again sponsor a Late-Breaker Epidemiologic Exchange at APHA's annual meeting in Chicago, Illinois. The Exchange, to be held on Wednesday, October 25, 1989, will provide a forum for presentation of epidemiologic investigations, studies, methods, etc , which have been conceived, conducted, and/or concluded so recently that abstracts could not meet the deadline for submission to other Epidemiology Sessions. Papers submitted should report on work conducted during the last 6–12 months.

Abstracts should be limited to 200 words; no special form is required. Abstracts should be submitted to Robert A. Gunn, MD, Division of Field Services, Epidemiology Program Office, Bldg. 1, Room 5127, C08, Centers for Disease Control, Atlanta, GA 30333—telephone; (404) 639-3187—and must be received by *September 15, 1989*.

The Reliability of Self-Reported Cigarette Consumption in the United States

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Abstract: To investigate the possibility that self-reported smoking is not a valid measure for assessing trends in smoking prevalence, we compared total self-reported cigarette consumption with the adjusted consumption data from cigarette excise taxes as reported by the US Department of Agriculture (USDA) for the period 1974 through 1985. Self-reported consumption was calculated by using data from the National Health Interview Surveys (NHIS) for adults and from the National Household Surveys on Drug Abuse for adolescents. For this period, the average ratio of self-reported cigarette consumption to the USDA estimate of consumption was 0.72 (range = 0.69 to 0.78). There was no statistical difference in this consumption ratio from year to year, indicating no apparent increase in the underreporting of cigarette smoking in these surveys. We conclude that cross-sectional surveys of self-reported smoking status remain a reliable surveillance tool for monitoring changes in population smoking behavior. (*Am J Public Health* 1989; 79:1020–1023.)

Introduction

Cigarette smoking has been identified as one of the major public health hazards of the 20th century. There are two standard approaches to measure the degree of exposure of the United States population to this hazard. The first approach uses information obtained in the process of collecting state and federal excise taxes and is an aggregate measure of population consumption. These data are regularly reported as total and per capita cigarette consumption by the Economic Research Service of the US Department of Agriculture (USDA). The advantage of the aggregate measure is that it is based on objective data that would not be affected by changing social attitudes towards smoking. The second approach is to survey a representative sample of the public and to ask questions regarding individual smoking behavior. This has been undertaken on a regular basis since the mid-1960s for the population of the United States by the National Center for Health Statistics and has been the basis of trend analyses of smoking reported elsewhere.[1–3] This latter measure, however, depends on the honesty and accuracy of respondents in reporting their own behavior.

The social acceptability of a behavior has been known to affect a person's willingness to admit engaging in that behavior[4] and may also affect the accuracy in reporting the number of cigarettes smoked per day. The social acceptability of smoking has decreased with decreasing smoking prevalence and increasing restrictions on smoking in public places.[5]

Using data from four national smoking surveys undertaken between 1964 and 1975, Warner concluded that the self-reported survey estimates of cigarette consumption increasingly underestimate consumption estimates from USDA during the period 1964 through 1975.[6] While he noted that this may have resulted from a change in survey methodology, he suggested that the finding was real and resulted from the increasing social stigma and personal health risk attached to smoking. If there has been increasing underreporting of smoking behavior, then the dramatic decrease in self-reported smoking prevalence reported elsewhere may be, in part, artifactual.[1,2]

Policy decisions with regard to smoking rely heavily on data from national surveys. It is essential therefore to document the validity and reliability of this method of measurement. In this report, we compared trends in the USDA estimates of cigarette consumption from 1974 to 1985 to estimates of consumption obtained from self-reported survey data collected during the same period by the National Center for Health Statistics and the National Institute on Drug Abuse. We also examined whether the trends in per capita consumption match the decline in smoking prevalence.

Methods

To calculate cigarette consumption, data from multiple sources were used. USDA publishes estimates of aggregate cigarette consumption. To estimate self-reported consumption, smoking behavior data from the National Health Interview Surveys and the National Household Surveys on Drug Abuse were analyzed. These data sources are described below.

USDA Estimate

The Economic Research Service of the US Department of Agriculture has made estimates of total and per capita consumption of cigarettes for a number of years. The estimates are based on data from the Bureau of Alcohol, Tobacco and Firearms of the US Department of Treasury, the Bureau of Commerce of the US Department of Commerce, the Tobacco Institute, and other private industry sources.[7]

The Tobacco Institute reports the number of packs of cigarettes on which state taxes are paid; the Bureau of Alcohol, Tobacco and Firearms reports the number of cigarettes on which federal taxes are paid; and the Bureau of Commerce reports the number of cigarettes imported into the United States. Both federal and state excise taxes are collected at the wholesale level (on removal from warehouses) and are not sales taxes. The tax-based data are then adjusted for estimated inventory changes and for returned unused products. Adult per capita consumption is calculated by dividing total consumption by the total US population 18 years of age and older for any given year.

Self-Reported Adult Consumption

The National Center for Health Statistics (NCHS), through the National Health Interview Survey (NHIS), has

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collected and made publicly available health information from a probability sample of the civilian, noninstitutionalized population of the United States since 1964. In cross-sectional surveys in 1974, 1976, 1978, 1979, 1980, and 1985, NCHS administered a supplemental questionnaire on tobacco use to a subsample of the population surveyed. These surveys included adults aged 18 years and older in 1983 and 1985, aged 19 years and older in 1976, and aged 17 years and older in 1974, 1978, 1979, and 1980. Detailed descriptions of the NHIS design and procedures have been published.[8,9] While the sample design of the NHIS changed in 1985, the changes were aimed at making the population estimates more precise. They do not invalidate comparisons among the estimates for the different survey years.

These supplemental surveys on tobacco use include self-respondents only (proxy responses were not allowed). When a selected respondent could not be interviewed in person, the survey was administered by telephone. The average response rate for the tobacco use supplements was 89.8 percent, with a range of 88.5 percent to 91.1 percent (NCHS unpublished data) and the sample size ranged from 9,553 in 1980 to 31,082 in 1985.[1]

The data were classified into age-sex-race groups and weighted to reflect the distribution of the US population for the survey year and the individual probability of selection as well as other survey design features. Ever smokers were defined as persons who reported that they had smoked at least 100 cigarettes in their lifetime. They were further identified as current smokers or former smokers based on responses to the question "Do you smoke now?" Respondents were asked how many cigarettes they smoked per day. We included in the analysis only survey data collected after 1974 because data collected prior to 1974 included proxy responses. Proxy reporting has been shown to underestimate the number of cigarettes smoked per day.[10]

Self-Reported Adolescent Consumption

Data on adolescents' smoking behavior are not available from the NCHS. However, the National Institute on Drug Abuse (NIDA) has conducted national household surveys of the civilian, noninstitutionalized population aged 12 years and older in the contiguous United States (Alaska and Hawaii excluded) in 1974, 1976, 1977, 1979, 1982 and 1985. These surveys use a self-administered questionnaire, administered in the household, and are based on a national probability sample of households. Respondents were asked the average number of cigarettes smoked during the past 30 days. The data are weighted to reflect selection probabilities based on age, sex, and race, and adjustments are made for nonresponses. The sample size in these surveys for those aged 12-17 years was approximately 2,000 for each survey year.

Analysis

Consumption Estimates—For each year, the total cigarette consumption is the amount smoked by: adult current smokers, adolescent current smokers, and former smokers who stopped smoking during that year. For each of these groups, the daily consumption is the product of the following three variables: 1) proportion of the group that is smoking (smoking prevalence); 2) the population size in each group; and 3) the average number of cigarettes smoked per day by each smoker in the group. Therefore, in each survey year, the consumption for each group equals the reported daily consumption times 365 days per year.

Calculations for adults were made using the public use data tapes available from NCHS for the years in which the supplements on smoking were conducted. Data for the adolescents were estimated from the NIDA surveys. We relied on published data for all years except 1985[11]; unpublished 1985 data were provided by NIDA (Joe Gfroerer, personal communication). Because the NIDA survey years differ from those of the NHIS, the prevalence and the average cigarette consumption were estimated by using linear intrapolation to match the NHIS years.

Census data were used to provide estimates of the appropriate age-specific population in the United States for each survey year.[12]

For current smokers covered by the NHIS, calculating the yearly consumption estimates was straightforward, using the smoking prevalence and the reported number of cigarettes smoked per day. Consumption by former smokers who quit in the year preceding the survey was estimated by using data from the combined 1978–1980 NHIS. For these three years, former smokers who quit during the previous year smoked an average of 192 days during that year, very close to half of the year. Accordingly, we assumed that former smokers who quit prior to the survey smoked for half of the year.

Because some adolescents also began to smoke during the year of the survey, self-reported cigarette consumption during the 30 days prior to the survey multiplied by 12 months overestimates adolescent yearly consumption. We had no valid and reliable national data to adjust for this overestimate. On the other hand, our data underestimate actual adolescent consumption because they do not include consumption by "experimental smokers" who had smoked during the year but had not done so in the 30 days prior to the survey. We conducted a sensitivity analysis to determine how assumptions made in calculating cigarette consumption affected the results of the analysis.

Consumption Ratios—Using both sources of cigarette consumption, we calculated a consumption ratio for each year by dividing the total self-reported cigarette consumption by the USDA estimate of consumption. This measure expresses the self-reported data as a fraction of the USDA estimate.

Statistical Analysis—Data were analyzed using STATISTIX,[13] an interactive program for microcomputers. A least squares line of best fit was calculated for both the per capita consumption and for the consumption ratios.

Results

The USDA estimate of cigarette consumption in the United States increased steadily from 1974 through 1981, when an estimated total of 640 billion cigarettes were smoked. Since 1981, this estimate of cigarette consumption has steadily declined despite increasing population size; in 1986, the number of cigarettes consumed was 584 billion (Table 1).

Per capita consumption decreased steadily from 4,141 cigarettes per year in 1974 to 3,274 in 1986 (Table 1). A regression line was fitted to these data and this line described the data well ($R^2 = 0.95$). The average reduction in per capita consumption was 75.4 cigarettes per year, (95% CI = 65.1, 85.7).

The USDA estimate of total cigarette consumption, self-reported cigarette consumption (based on NCHS and NIDA surveys), and the consumption ratios for the selected years from 1974 through 1985 are shown in Table 2. Although self-reported consumption is less than the USDA estimate of

TABLE 1—USDA Estimates of Total Cigarette Consumption and per Capita Consumption, United States, 1974 through 1986

Year	Estimated Total Cigarettes Consumed per Year (billions)	Estimated* per Capita Consumption of Cigarettes Per Year
1974	599.0	4,141
1975	607.2	4,123
1976	613.5	4,092
1977	617.0	4,051
1978	616.0	3,967
1979	621.5	3,861
1980	631.5	3,844
1981	640.0	3,836
1982	634.0	3,739
1983	600.0	3,488
1984	600.4	3,446
1985	594.0	3,370
1986	583.8	3,274

*Based on US population 18 years of age and older; manufactured cigarettes only.
SOURCE: US Department of Agriculture.

consumption at each year, this relationship appears to be constant over time. The consumption ratios ranged from 0.78 in 1983 to 0.69 in 1976, with a mean of 0.72. The slope of regression line is not significantly different from zero ($p = 0.85$), strongly suggesting the lack of any change over time in the consumption ratio. The components of the self-reported cigarette consumption data for each year and for each of the three groups are shown in Table 3. The contributions of adult former smokers and adolescent smokers are small relative to the amount smoked by adult current smokers.

Sensitivity Analysis

For the former-smoker and the adolescent components of total self-reported consumption, assumptions that were used in calculating self-reported consumption were:

- former smokers who quit during the preceding 12 months quit on average at the mid-point of the year; therefore, they were smoking for 182 days (rather than 192 days estimated from the 1978–80 surveys);
- adolescent data could be interpolated from the NIDA survey years to the NHIS years;
- the impact of adolescents who took up smoking during the preceding year and the impact of the experimental smokers who did not smoke in the 30 days prior to the survey were small and canceled each other out.

The results were not altered when 192 days was used instead of 182 days. The mean consumption ratio was

TABLE 2—USDA and Self-Reported Cigarette Consumption Estimates, United States, Selected Years 1974 through 1985

Year	USDA Estimate (billions)	Estimate from Self-Reported Data (billions)	Consumption Ratio
1974	599.0	434.86	0.726
1976	613.5	424.36	0.692
1978	616.0	438.38	0.712
1979	621.5	441.20	0.710
1980	631.5	459.11	0.727
1983	600.0	467.83	0.780
1985	594.0	414.35	0.698

SOURCES: US Department of Agriculture, National Center for Health Statistics (National Health Interview Survey), National Institute on Drug Abuse (National Household Survey on Drug Abuse).

TABLE 3—Self-Reported Cigarette Consumption among Different Population Sub-groups, United States, Selected Years, 1974 through 1985

	Estimated Cigarette Consumption in Billions by			
Year	Adult Current Smokers	Adult Former Smokers*	Adolescent Current Smokers	Total Population
1974	393.97	21.38	19.51	434.86
1976	387.34	20.30	16.72	424.36
1978	404.02	19.47	14.89	438.38
1979	405.59	22.44	13.17	441.20
1980	423.80	23.02	12.29	459.11
1983	432.43	25.02	10.38	467.83
1985	377.82	26.26	10.26	414.35

*Former smokers who had quit smoking during the preceding year.
SOURCES: National Center for Health Statistics (National Health Interview Survey), National Institute on Drug Abuse (National Household Survey on Drug Abuse).

affected at the third decimal place (0.720 versus 0.721). This lack of effect is partially due to the small contribution of former smokers to the total amount of the self-reported consumption. The contribution of the adolescent smoking component to total consumption was even smaller than that of former smokers for each survey year (Table 3). It ranges from 2.2 percent to 4.5 percent of the total self-reported consumption. If it would increase by 50 percent at each year, the mean consumption ratio would be affected at the second decimal place (0.720 vs 0.732); therefore variations in this component are not likely to influence the results substantially either. Given this small impact of adolescent smoking, an overlap between the two surveys for the lower age bound (those aged 17)—which occurred in the years 1974, 1978, 1979, and 1980—is not likely to have a major effect. Overall, over 90 percent of the total self-reported consumption is attributed to current smokers for whom relevant data are readily available.

Discussion

Measurement of any quantity raises issues of validity and reliability. The findings of this analysis, comparing USDA estimates of cigarette consumption to self-reported smoking consumption, 1974 through 1985, address both questions. The consumption ratio 0.72 represents a considerable discrepancy between the two measures. It is unlikely that this discrepancy between the USDA and the self-report estimates stems from the former and adolescent components of the total consumption because they account for less than 10 percent of the total consumption. It may be that adult current smokers do not report their smoking status correctly, but the literature does not support this conclusion. Findings from biochemical validation studies[14,15] suggest that self-reported cigarette consumption prevalence is a valid estimate of smoking status in the population. An alternative explanation may be that current smokers consistently underreport the number of cigarettes they smoke. This underreporting may occur disproportionately among those who are heavier smokers, analogous to the underreporting of alcohol intake.[16]

Alternatively, it might be from a systematic rounding down of the number of cigarettes smoked per day as respondents have a preference to report daily consumption in multiples of a half-pack of cigarettes and therefore report the closest half-pack consumption below their actual consumption level.[5,17]

The ratio between the two estimates is constant over time, strongly suggesting that self-reported data from these national surveys are reliable, despite any increasing social unacceptability of smoking which could theoretically bias the accuracy of self-reported data. A bias could occur if the two components of the self-reported consumption estimate, i.e., prevalence and daily consumption, change differentially over time. Prevalence has been decreasing steadily since 1974.[1,2] Daily consumption, however, remained relatively constant over time.[5]

These findings differ from the conclusions drawn from the study by Warner[6] who reported that the consumption ratio decreased from 0.73 in 1964 to 0.64 in 1975. This earlier analysis used data from the 1964, 1966, 1970, and 1975 Adult Use of Tobacco Surveys. These surveys, however, used different methodologies, particularly with respect to the mode of data collection. The 1964 and 1966 surveys used in-person interviews, whereas the 1970 and 1975 surveys used telephone. The consumption ratios obtained for 1964 and 1966 (0.73 and 0.72, respectively), using in-person survey data, are similar to the mean ratio of 0.72 observed in our study from in-person surveys during the period 1974 to 1985. Therefore, there is no evidence for an increase in underreporting when comparing in-person survey data from the 1960s with data from 1974 to 1985. The drop in consumption ratio reported by Warner between 1966 and 1970 is probably best explained by the change in the interview mode, because telephone surveys generally provide slightly lower estimates than in-person household surveys.[5]

One might ask whether it is more appropriate to use the USDA estimates of cigarette consumption, which are more objective and easily available, or self-reported data such as the NCHS or the NIDA surveys, as the primary source for surveillance of smoking in the population. The USDA estimate is a single overall figure. However, such a figure does not provide the number and types of people who smoke or the number of cigarettes each smoker consumes and hence cannot be used to target antismoking programs. On the other hand, trend analyses of self-reported smoking[1–3] do provide the smoking behavior of individuals in the population (and in sociodemographic subgroups) both in terms of smoking status and daily consumption patterns. The aggregate estimate may be more sensitive to influences of anti-smoking activity than estimates of smoking prevalence because, for many people, it may be easier to reduce the number of cigarettes smoked per day than to quit smoking. This aggregate estimate, however, cannot distinguish between a large number of people reducing the amount they smoke and a smaller number who actually quit.

The most important public health goal is to reduce smoking prevalence, not simply to reduce the number of cigarettes consumed by each smoker. Even though a reduction in daily cigarette consumption has some health benefits, these are smaller than those of quitting smoking.[5,18] Our results support the use of self-reported cigarette smoking from national surveys as a reliable surveillance tool, essential in monitoring progress toward achieving the 1990 health Objectives for the Nation,[19] and a smoke-free society by the year 2000.

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Two Episodes of Acute Illness in a Machine Shop

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Abstract: Following an explosion in a machine shop and temporary plant closure, on the day the plant returned to full operations a degreaser malfunctioned. Workers in the assembly room were exposed to trichloroethylene levels later estimated to have exceeded 220 ppm (OSHA PEL 100 ppm). The plant was evacuated and the degreaser taken out of operation. Blood testing for carbon monoxide (CO) on five employees found carboxyhemoglobin levels in excess of normal. The plant reopened the following morning. Over the next two weeks, 15 employees were seen by the plant nurses for similar complaints; although all returned to work, their carboxyhemoglobin levels, later found to be inaccurate, were reported by a local medical clinic to range from 13.7 to 20.0 percent. At the end of the second week, another outbreak of illness occurred, but carboxyhemoglobin, trichloroethylene, fluorocarbons, and methylene chloride were not elevated in all 17 persons tested; plant-wide monitoring for CO found no elevated levels. During the first outbreak of illness, cases were 2.26 times as likely to have entered the assembly room as noncases. During the second outbreak, cases were no more likely than noncases to have entered the assembly room. We believe the explosion, earlier toxic exposures and illness, and the misleading blood test results led to plant-wide anxiety which culminated in a collective stress reaction and the second outbreak. An open meeting with all employees, informing them of our findings, provided reassurance and no further episodes of illness occurred in this workforce. (*Am J Public Health* 1989; 79:1024–1028.)

Introduction

Disease is the result of the interaction between a host, an agent, and the environment. The host must be susceptible to the agent and the environment dictates if the agent and host contact each other. In this paper, we describe two outbreaks of illness in a manufacturing plant within 16 days. Whereas the workers' symptoms were similar on both days, the agents and environments differed. Although the outbreaks were not independent of one another, we believe they represented separate illnesses with unique causes.[1]

The 225,000 sq ft factory manufactures rotary compressors for small air conditioners. At the facility, internal compressor parts are machined to specifications using a variety of machine tools. Welding, brazing, painting, and soldering operations are used as appropriate during compressor assembly. The majority of the workers are unskilled and assigned to the same daily task with little or no variation in their day-to-day routine. A floor plan of the plant is presented in Figure 1.

Chronology of Events

A time line and chronology of events is presented in Figure 2.

On January 29, 1986, a regulator valve on the fuel line to the plant boiler malfunctioned resulting in a large explosion. Fortunately, this occurred while second shift workers were on break in a second floor lunchroom. No one was injured.

The plant reopened on February 3. First shift began at 7:30 am. Between 8:35 am and 12:30 pm, 10 employees became acutely ill and were taken to the local emergency room where they were evaluated at a local hospital for a presumed toxic inhalation. Their symptoms included headache, dizziness, disorientation, eye and throat irritation, nausea, vomiting, shortness of breath, and chest pain. Arterial blood samples, drawn for carboxyhemoglobin determination, were elevated in five workers (mean 14.0 percent, range 13.3–14.8 percent). All workers were treated and released. By early afternoon, first shift production was stopped; second shift workers were sent home as they reported for work.

State OSHA (Occupational Safety and Health Administration) inspectors and an industrial hygiene consultant hired by the company arrived that afternoon. Direct-reading mea-

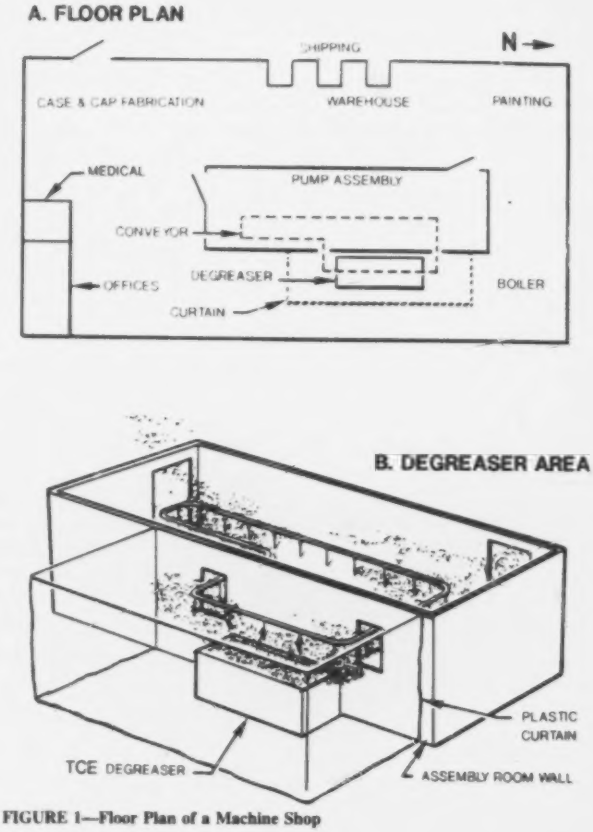

FIGURE 1—Floor Plan of a Machine Shop

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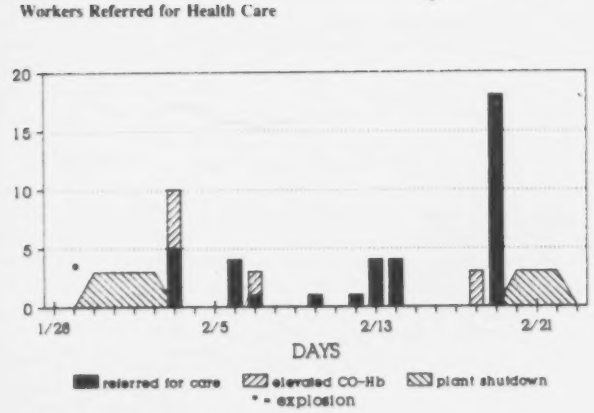


FIGURE 2—Time Line and Sequence of Events from Explosion January 28 through February 21

surements for carbonmonoxide (CO) were made throughout the plant but none exceeded the federal OSHA standard, the ACGIH threshold limit value (TLV), or the NIOSH (National Institute for Occupational Safety and Health) recommended standard.[2–5] Propane forklifts, open flames in the welding and brazing operations (located within the assembly room), and a portable heater were identified as CO sources. Two of seven forklifts, emitting excessive CO, were taken out of service, pending repair.

Workers interviewed the same day reported that the degreaser, started two hours before the beginning of first shift, had emitted trichloroethylene fumes into the assembly room. Following the explosion, the degreaser had been surrounded by polyurethane curtains to prevent contamination of compressor parts with residue from the explosion (Figure 1B). The curtains had enclosed not only the degreaser but also two openings in the assembly room wall through which parts were normally conveyed to and from the degreaser. On the morning of February 3, the degreaser malfunctioned. Apparently, the trichloroethylene level had fallen below the operating range and then overheated; excess fumes trapped within the polyurethane curtains escaped into the assembly room.

The degreaser was removed from operation but was restarted two days later to simulate the conditions of February 3. Measurements taken two hours after start up found trichloroethylene concentrations of 224 parts per million (ppm) in the assembly room. The OSHA permissible exposure limit (PEL) for trichloroethylene is 100 ppm as an 8-hour time-weighted average (TWA) and 200 ppm as a ceiling concentration.

Full production resumed the next day. Between February 4 and 18, 15 employees were seen in the plant infirmary with complaints of headache, eye and throat irritation, nausea, fatigue, and muscle aches. Twelve employees sought treatment at local medical facilities because of concern about continuing exposures at work. One clinic drew venous blood samples for carboxyhemoglobin determination from five workers. All were elevated, ranging from 13.7 to 20 percent. Three workers were told of their elevated carboxyhemoglobin levels while at work, during the morning of February 19. Local newspaper accounts of the February 3 mishap were published daily over the next four days. On February 13, the press had reported that another employee had been rushed to the hospital, "overcome by toxic fumes." All environmental samples for CO were well below the NIOSH recommended exposure level (REL) of 35 ppm.

On February 19, between 9 am and 11 am, workers again became acutely ill. Most were triaged from the single room plant infirmary where they could be found sitting or lying on either the floor, examining table, or stretchers. Two workers lost consciousness for a brief period. Fifteen workers were transferred by ambulance to the hospital. Emergency medical technicians carried stretchers through the plant production area as they left the facility. Four additional workers later requested a medical evaluation. Production at the plant was stopped.

At the hospital, 17 employees were tested for elevated carboxyhemoglobin levels, all tested normal (mean 2.3 percent, range 0.0–8.0 percent). Additional toxicologic screening for trichloroethylene, fluorocarbons, and methylene chloride was also negative.

State OSHA investigators inside the plant during the outbreak found CO levels less than 15 ppm.

NIOSH Investigation

NIOSH investigators arrived the next day. Our evaluation included an environmental survey and a review of the medical findings. With the exception of the trichloroethylene exposures on February 3, no environmental sources of toxic exposures likely to have caused acute illness were identified.

A review of medical records showed that treatment had consisted of monitoring vital signs, administering oxygen, and an occasional precautionary intervenous line. Although CO levels inside the plant were low, all carboxyhemoglobin levels in employees seen at the clinic from February 4–18 were elevated. At the same time, no carboxyhemoglobin levels were elevated at the hospital after February 3. One of us (PRK) submitted samples of his blood for carboxyhemoglobin determination on February 20. Results from the hospital were 0.03 percent, while the clinic results were 16.0 percent. Two additional samples, collected at the clinic from employees on February 21 (two days after the plant was closed) were also elevated, with carboxyhemoglobin levels of 16 percent and 20 percent. The inaccuracy of the clinic carboxyhemoglobin values were later confirmed by the regional laboratory director.

During meetings with the union and management, the decision was made to re-open the plant. An open meeting was held for all employees on February 23 informing them of our findings. NIOSH investigators reassured workers that the plant environment was safe and would continue to be monitored. Plant operation began immediately following the morning meeting.

Questionnaire

On February 25, a questionnaire was administered to 325 employees (279 first shift and 46 second shift workers). Second shift workers and eight others who had not worked on February 3 and February 19 or did not adequately complete the questionnaire were excluded. Employees ranged in age from 18 to 62 years with a mean of 33 years; 67 percent were

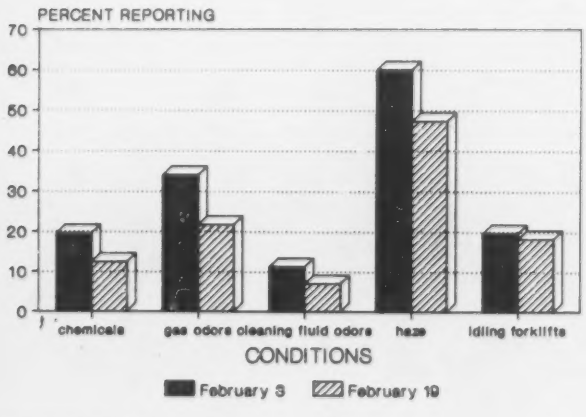
male. Almost 30 percent of the employees had not finished high school, 42 percent were high school graduates, 21 percent had some college or vocational training, and 7 percent were college graduates.

Figure 3 illustrates the proportion of employees who reported the presence of certain conditions in the plant. The presence of chemicals, odors of gasoline and cleaning fluid, and visible haze was reported by a greater percentage of workers on February 3 than on February 19.

Headache was the most common symptom on both February 3 and 19, with greater than 20 percent of the workforce reporting it. Other reported symptoms included sleepiness, sore throat, lightheadedness, nausea, and weakness. The percentage of employees reporting each symptom was similar on the two days. In order to define a case, we used the occurrence of headache plus a varied number of additional symptoms. If the number of symptoms experienced by an individual had been a function of exposure, we would expect that workers with an increased number of symptoms had received an increased exposure. This would result in a case definition that became more specific for exposure as it became more restrictive.

Employees were asked to recall their time spent in the assembly room on February 3 and February 19. Eighty-seven of 240 and 79 of 245 employees entered the assembly room on February 3 and February 19, respectively. On February 3, entry into the assembly room was associated with illness; the relative risk ranged from 1.33 to 3.10, increasing as the case definition became more restrictive (Table 1). No association between illness and entry into the assembly room was observed on February 19. Time spent in the assembly room was also associated with illness on February 3, but not on February 19 (Table 2).

On both days, women were more likely than men to become ill. On February 19, women throughout the plant had a two-fold increased risk of illness (RR = 1.99, 95% CI = 0.93, 4.23). In contrast, on February 3, women entering the assembly room were at greatest risk (RR = 3.15, 95% CI = 1.06, 9.37) while the RR for women versus men who did not enter the assembly room on February 3 was much lower (RR = 1.24, 95% CI = 0.35, 4.43).



Reported by 1st shift workers

FIGURE 3—Workers' Perceptions of Conditions Inside a Machine Shop during Two Episodes of Acute Illness.

Discussion

Our findings suggest that the cause of illness on February 3 differed from that on February 19. The illness of February 3 appears to have been the result of the toxic effects of CO complicated by exposure to trichloroethylene. The illness of February 19, however, more likely represents a collective stress reaction. The combination of the boiler explosion on January 29, toxic exposures on February 3, reports in the local press, and the misleading laboratory results reported to employees on February 19 led to plant-wide anxiety, which precipitated the acute outbreak of illness on February 19.

On February 3, presence in the assembly room was associated with illness and this association was strengthened by both an increase in the number of symptoms and time spent in the assembly room. Unfortunately, the cause of the elevated carboxyhemoglobin levels of February 3 was not identified. Contact with strong alkalies may cause trichloroethylene to decompose into CO, phosgene, and dichloroethylene.[6] However, no strong alkalies were found in the assembly room and symptoms suggestive of phosgene poisoning were not reported.

The February 19 outbreak shares a number of common features with other reports of collective stress reaction.[7–12] No exposures were identified in the work environment at the time of the illness. The ill workers developed a wide variety of relatively severe symptoms, which contrasted sharply with a lack of physical findings. The spread of illness was very rapid, resembling an acute chemical poisoning rather than an outbreak of infectious disease. Perhaps most important, a very high level of stress had been building in this workforce since the boiler explosion on January 29.

In the workplace, collective stress reaction has been reported most often among groups of unskilled workers with less than a high school education performing highly repetitive unskilled tasks.[11] In those reports, women were more likely to have been affected than men. It should be emphasized that the association of gender and educational level with collective anxiety phenomenon does not imply that these are causative factors. The nature of the job, rather than gender or education, may place these workers at high risk.

Whenever workers are exposed to toxic substances, both they and their unexposed co-workers are subjected to psychological stress and the situation may produce collective stress reactions above and beyond the primary symptoms. This, in turn, can precipitate additional illness episodes, such as the one illustrated in this paper. This emphasizes the importance of a thorough and sensitive debriefing of the workers following any environmental or medical assessment. If worker fears and suspicions are not sufficiently allayed and episodes of illness do recur, both the investigators and the management can lose credibility, and the situation may escalate.

The conclusion that an outbreak is the result of collective stress reaction should be made with caution. There is always the chance that an undetected toxic exposure has been overlooked.[13–15] Illness resulting from psychologic factors is "real" in every respect. The term refers to illness in which the primary cause is psychological stress, arising from the occupational and/or general social environment, rather than environmental chemical, physical, or infectious agents. The occurrence of a collective stress reaction does not imply a psychiatric disorder. It represents a normal psychophysiological response to a stressful environment.

TABLE 1—Percentage of Ill Workers by Exposure in the Assembly Room, Risk Ratios, and p-values

	Headache + 1 Symptom	Headache + 2 Symptoms	Headache + 3 Symptoms	Headache + 4 Symptoms	Headache + 5 Symptoms
February 3					
% Assembly Room employees ill (N = 87)	17.2	16.2	13.1	13.1	12.1
% Other employees ill (N = 153)	13.0	9.7	5.8	5.2	3.9
Risk Ratio	1.33	1.68	2.17	2.53	3.07
95% CI	0.75,2.36	0.90,3.12	1.03,4.55	1.18,5.42	1.33,7.04
February 19					
% Assembly Room employees ill (N = 79)	15.4	12.1	10.3	8.8	8.8
Other employees ill (N = 166)	13.1	10.9	9.7	8.0	4.6
Risk Ratio	1.16	1.15	1.03	1.09	1.97
95% CI	0.63,2.16	0.58,2.27	0.49,2.20	0.48,2.47	0.80,4.88

*X^2 test

TABLE 2—Illness by Duration of Exposure in the Assembly Room

Case Definition	% ill *(RR)	% ill (RR)	% ill (RR)	Summary RR	95% CI
Minutes in Assembly Room:	0	1–60	61–390		
		February 3, 1986			
Headache + 1 symptom	13.07	13.9	19.6		
	(1.00)	(1.06)	(1.41)	1.27	0.79,2.05
Headache + 2 symptoms	9.8	11.1	19.6		
	(1.00)	(1.12)	(1.75)	1.50	0.93,2.42
Headache + 3 symptoms	5.9	5.6	17.6		
	(1.00)	(0.95)	(2.21)	1.77	1.07,2.95
Headache + 4 symptoms	5.2	5.6	17.6		
	(1.00)	(1.05)	(2.36)	1.91	1.16,3.15
Headache + 5 symptoms	3.9	2.8	17.6		
	(1.00)	(0.74)	(2.70)	2.12	1.28,3.52
Number of employees	153	36	51		
		February 19, 1986			
Headache + 1 symptom	13.9	17.2	12.0		
	(1.00)	(1.24)	(0.88)	1.01	0.57,1.79
Headache + 2 symptoms	11.4	17.2	8.0		
	(1.00)	(1.48)	(0.73)	1.01	0.54,1.90
Headache + 3 symptoms	10.2	13.8	6.0		
	(1.00)	(1.33)	(0.63)	0.88	0.44,1.82
Headache + 4 symptoms	8.4	10.3	6.0		
	(1.00)	(1.21)	(0.75)	0.92	0.43,1.96
Headache + 5 symptoms	4.8	10.3	6.0		
	(1.00)	(1.93)	(1.19)	1.47	0.72,3.00
Numbers of employees	166	29	50		

*RR = the risk among the exposed/the risk among the unexposed.
†Mantel-Haenszel X^2 test for dose-response.[15]

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The 1990 Census: Size and Scope

The 1990 Census of the United States will be the 21st in the nation's history and the largest and most complex ever undertaken. The Census Bureau has been planning the mammoth operation since 1984. The size and scope of the 1990 Census will include:

- *Expected count*—250 million people, 106 million housing units.
- *Employment*—Will hire about 565,000 people, with 300,000 working at peak.
- *Location:* 484 field offices, 13 Regional Census Center, 7 processing offices.
- *Coverage*—Over 106 million questionnaire packages to print, label, and process.
- *Geography*—Computerizing entire map base, including over 250,000 different base maps. Nearly 7 million maps will be generated to help census takers collect questionnaires.
- *Participation*—Over 70 million mail returns within two weeks of Census Day, April 1, 1990.
- *Technology*—Using 570 minicomputers in field offices, to be set up, used, and dismantled in one year.
- *Mandated Completion*—Deliver apportionment counts to the President by December 31, 1990; provide data necessary for redistricting to states no later than April 1, 1991.

Most households will receive a mailed questionnaire about a week before Census Day. Others, depending on location, will have forms delivered by census takers. Roughly five out of six households will receive a short form; one in six will be asked to fill out a longer form.

The questions are basic. The *short form* asks about race, Hispanic origin, age, marital status, rent or own residence, number of rooms, etc. The *longer form* asked additional questions on ancestry, employment, education, income, type of housing, utilities, etc.

The population count for various areas of the nation is put to numerous uses, e.g., to determine how many seats a state will have in the US House of Representatives; to allocate seats in state legislatures, and to benefit specific populations and programs. Billions of dollars of federal funds are returned to states, local governments, and to American Indian reservations and Alaska Native villages. State funds are returned to cities and counties, with amounts based in part on population or housing data. Planners use census facts to determine where to locate hospitals, schools, day care and senior citizen centers. Various special programs receive funds based on census information. Businesses use the data for decision-making involving expansion and jobs.

For a number of reasons, about 2.2 million people were not counted in the last US Census (about 1 percent). Among minority populations, the undercount was higher—about 6 percent. To help assure a complete count in the 1990 Census, the Bureau of the Census began distributing information in June to US local and tribal government officials. Known as the Local Review Program, the effort encourages officials in the nation's 39,000 local and tribal governments to voluntarily check Census Bureau maps and block-by-block housing unit counts compiled from a master list created and updated throughout the census period. Coupled with that mailout, approximately 470 half-day local review training sessions will be held throughout the US before the end of 1989.

For further information about the Local Review Program and other Census matters, contact: Chief, Decennial Planning Division, Bureau of the Census, Washington, DC 20233. Tel: Census Promotion Office, Information Service, Census '90, (301) 763-1990. Or, contact Census Bureau Regional Census Centers in: Atlanta, GA, Boston, MA, Charlotte, NC, Chicago, IL, Dallas, TX, Denver, CO, Detroit, MI, Kansas City, MO, Los Angeles, CA, New York, NY, Philadelphia, PA, San Francisco, CA, and Seattle, WA.

Lead Absorption in Indoor Firing Range Users

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Abstract: To determine if users of indoor firing ranges may be at risk from lead exposure, we studied a law enforcement trainee class during three months of firearms instruction. Blood lead levels were obtained before training and at four-week intervals during training. Air lead levels were measured three times during instruction. Blood lead levels rose from a pre-training mean of 0.31 μmol/L to 2.47 μmol/L. Mean air lead levels were above 2,000 μg/m^3, more than 40 times the Occupational Safety and Health Administration's standard of 50 μg/m^3. Cumulative exposure to lead and the change in blood lead were positively correlated. Control measures need to be studied to determine their efficacy in decreasing or eliminating this health risk. (*Am J Public Health* 1989; 79:1029–1032.)

Introduction

Indoor firing ranges have been the subject of several studies of occupational lead toxicity. These investigations have documented elevated blood lead levels and associated adverse health effects in employees and instructors at these ranges.[1–5]

Lead exposure in indoor firing ranges occurs primarily through inhalation of lead particulates suspended in the range air. The major sources of airborne lead in the breathing area of the shooter result from the ignition of primer material containing lead styphnate (a highly explosive compound used to initiate the combustion of gunpowder in the cartridge) and, because the gun barrel and the bullet do not always align exactly, the shearing of lead particulates off of the bullet as it passes through the weapon. Fragmentation of bullets when they strike the target or backstop may also contribute to the overall air lead concentration in the firing range.[1,6–8]

In 1985, the Colorado Department of Health began an Occupational and Environmental Disease Surveillance Program monitoring exposure to heavy metals as well as other environmental exposures and work-related outcomes. Through this program, laboratories were requested to report elevated values for lead, cadmium, arsenic, and mercury. (Since October 1988, elevated blood or urine levels for lead, cadmium, arsenic, and mercury have been required to be reported to the Colorado Department of Health.) During 1985 and 1986, 11 reports of elevated blood lead levels were received; about half from employees of indoor firing ranges and the rest from frequent users of indoor firing ranges. These persons had blood lead levels determined for various reasons; awareness that use of an indoor firing range may be a health hazard, physician impression of lead toxicity, and requirement for an employment physical. While occupational studies have suggested that exposure to lead from indoor firing ranges may be a health risk for frequent users of these ranges,[2,6,9] the extent of risk among users has not been documented. This concern prompted us to conduct a study to determine the amount of lead absorbed into the body by regular users of an indoor firing range.

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Methods

Seventeen members of a beginning law enforcement trainee class were enrolled in this study on January 29, 1987, before the start of any firearm training or practice on the range. After informed consent was obtained, questionnaires to obtain demographic information and to determine any other potential sources of lead exposure were completed and blood was drawn to obtain baseline values for blood lead and free erythrocyte protoporphyrin (FEP) levels. A second questionnaire was administered on May 3, 1987, after firearms training was complete, to assess any self-reported symptoms of lead toxicity and to again ascertain any other potential sources of lead exposure which may have occurred during the training period.

The class began firearms instruction on February 3, 1987, and ended on April 28, 1987. A log was kept by the trainees indicating the booth each person occupied and the time spent on the range for each day of training. Smith and Wesson* stainless steel K frame Model 66 .357 caliber revolvers and .38 special caliber + P, reloaded, lead semi-wad cutter cartridges, manufactured by 3D INV* were used throughout the firearms training.

Blood was drawn at approximately four-week intervals (January 29, March 3, March 31, and May 3) to determine blood lead and FEP levels. After training was complete, the class dispersed throughout the state; only nine of the 17 trainees were available for blood sampling six weeks after completion of training. No physical examinations or other medical tests were done. Laboratory analyses were performed by ESA Laboratories in Bedford Massachusetts, using anodic stripping voltammetry without digestion for blood lead levels and the Chisolm Brown extraction method for FEP levels.[10–13] Since 0.24 μmol/L was the lower level of detection for blood lead, results that were <0.24 μmol/L on January 29 were set at 0.12 μmol/L for statistical analyses.

Air samples were taken from each occupied booth on the firing range on February 3, March 3, and March 31. Air sampling was also done in the classroom directly across the hall from the firing range and in the control room to determine exposure levels while in either room. Battery powered SKC model 224-37 universal flow personal air sampling pumps coupled with 37-mm mixed cellulose ester filters and closed cassettes were used for air sampling. The pump flow rate was set at approximately 2 liters per minute. Sampling pumps were calibrated before and after every sampling period using

*Throughout this paper, any mention of commercial names or products is for identification purposes only and does not constitute endorsement by the authors or the agencies for which they work.

a 2-liter bubble buret. Filters were analyzed by the Colorado Department of Health Laboratory using atomic absorption spectrophotometry. Field blanks were submitted with each set of samples for quality control. Pumps in the classroom and the control booth were placed on counter tops. The pumps on the range were attached to the overhead structure about six inches above the trainee's head. They were turned on when the trainees entered the range and turned off after their shooting session was completed.

During this study, two changes were made to the ventilation system of the firing range. The first, on February 5, corrected the positive pressure inside the range that had allowed lead-contaminated air to flow from the range into other parts of the building whenever the range door was opened. On March 9, final adjustments consisted of placing fins on the air supply grill to cause smoother air flow across the firing line and to decrease air turbulence. Air lead levels were determined before and after adjustments to the ventilation system. The air lead levels on February 3 were used as exposures for February 3 and 4 only. Air lead levels measured on March 3, were used as estimates of exposure levels from February 5 through the day before final adjustments were made. Results from the third air sample on March 31 were used as exposure levels after changes to the ventilation system were complete. In addition to the regular training, the March 31 session also included a very short (3 to 12 minutes) session of "dual-image" shooting with only one recruit on the range at a time. The air lead concentration determined for this session was used as the approximate exposure for the five times "dual-image" shooting was conducted.

Data were categorized into time periods based on the dates of blood sampling and firearms training: *Period 1*, January 29–March 2; *Period 2*, March 3–March 30; and *Period 3*, March 31–May 3. Each person's approximate exposure to airborne lead was established for each time period based on time spent on the firing range, the booth each person occupied, and air lead concentrations for each booth as determined from the three air samplings. For Period 1, we examined the change in the blood lead levels by the cumulative exposure to airborne lead. Cumulative exposure was obtained by multiplying the air lead concentration at each booth by the time spent on the range for each trainee each day of training and adding the totals for the six training days during this time period.

To assess the effect of using jacketed ammunition on the air lead concentrations, a final air sampling was done on April 30 using three different types of ammunition. Each test period consisted of the firearm instructors firing 60 rounds in approximately 10 minutes. In the first sampling period, copper-jacketed bullets were used; in the second, nylon-clad bullets; and in the third, the lead bullets that were normally used.

Results

Fifteen of the 17 trainees (ages 24 to 40, mean 30.7) were males. Only one person had a history of potential occupational exposure to lead during the previous year. His blood lead level upon entry into the study was the highest of the group, at 1.11 μmol/L. Baseline blood lead levels for the others were at or below 0.53 μmol/L, with a mean for these 16 of 0.26 μg/dl and a mean for all 17 of 0.31 μmol/L (Table 1).

TABLE 1—Blood Lead Levels of Law Enforcement Trainees, μmol/L, Colorado, 1987

Trainee	Date Sample Was Taken				
	1/29/87	3/03/87	3/31/87	5/03/87	(6/10–6/18)
1	0.53	1.73	1.59	1.59	1.01
2	0.29	2.32	2.46	1.93	
3	0.48	2.61	1.30	1.11	0.58
4	0.34	2.61	2.12	1.74	
5	<0.24	2.87	1.98	1.59	1.16
6	0.29	2.36	2.75	2.46	1.26
7	<0.24	2.99	2.51	2.08	
8	0.29	3.19	2.61	2.46	1.74
9	<0.24	1.50	1.69	1.79	
10	1.11	1.93	2.04	2.32	
11	0.29	2.79	1.93	2.03	
12	<0.24	2.03	1.59	1.25	
13	0.29	no data	no data	1.59	1.16
14	<0.24	3.52	2.65	2.22	1.45
15	0.29	1.98	1.88	1.79	0.97
16	0.24	2.37	2.12	2.12	
17	0.29	2.75	2.99	2.32	1.45
Mean	0.31	2.47	2.14	1.91	1.20
Median	0.29	2.49	2.08	1.93	1.16

Mean change 1/29–3/03 = 2.16 μmol/L
Mean change 3/03–3/31 = −0.34 μmol/L
Mean change 3/31–5/03 = −0.23 μmol/L

During firearms training, the class was on the range a total of 17 days: six days in *Period 1*; five in *Period 2*; and six in *Period 3*. The time spent on the range each day varied, depending upon the type of training being done. In addition, the number of rounds fired by each trainee each day also varied, but was approximately 120 rounds per hour of training (Table 2).

At the end of Period 1, after just seven hours of shooting over that four-week period, the blood lead levels were elevated, from a pre-training mean of 0.31 μmol/L to 2.47 μmol/L (Table 1). Thereafter the levels decreased slightly. The blood lead levels for all nine trainees who had a fifth blood sample drawn decreased between the fourth and fifth sample, from a mean of 1.90 μmol/L to 1.20 μmol/L.

The FEP levels rose throughout the entire study period from a pre-training mean of 0.49 μmol/L to 1.27 μmol/L at the end of the study (Table 3). The FEP levels for the nine trainees who had blood drawn about six weeks after completion of training all showed a decrease from the May 3 level, dropping from a mean of 1.53 μmol/L to 1.23 μmol/L.

Air lead concentrations for each occupied booth are

TABLE 2—Amount of Time the Law Enforcement Trainees Spent in the Firing Range and the Approximate Number of Rounds Fired, per Time Period

Time Period	Number of Days	Mean Total Time (hours)	Range	Mean Time per Session (minutes)	Range	Rounds Fired
1. (February 3–March 2)	6	7.2	7.0–7.6	72	33–120	900
2. (March 3–March 30)	5	2.6	2.5–2.7	39	12–78	300
3. (March 31–May 3)	6	4.1	4.1–4.1	41	3–121	550

TABLE 3—Free Erythrocyte Protoporphyrin Levels of Law Enforcement Trainees, μmol/L, Colorado, 1987

	Date Sample Was Taken				
Trainee	1/29/87	3/03/87	3/31/87	5/03/87	(6/10–6/18)
1	0.35	0.46	0.37	0.44	0.37
2	0.60	1.43	2.53	2.21	
3	0.62	0.66	0.92	1.12	0.90
4	0.34	0.48	0.62	0.64	
5	0.39	0.97	1.91	2.11	1.74
6	0.62	0.81	1.08	1.24	1.01
7	0.43	0.48	0.92	1.03	
8	0.37	0.51	0.81	0.94	0.87
9	0.35	0.37	0.48	0.48	
10	0.60	0.62	0.96	1.04	
11	0.35	0.50	0.94	1.35	
12	0.55	0.58	0.89	0.69	
13	0.76	no data	no data	1.84	1.56
14	0.67	1.04	2.16	2.66	1.93
15	0.46	0.66	0.76	0.80	0.64
16	0.34	0.35	0.43	0.48	
17	0.48	1.29	2.81	2.60	2.07
Mean	0.49	0.70	1.16	1.27	1.23
Median	0.46	0.60	0.92	1.00	1.01

Mean change 1/29–3/03 = 0.21 μmol/L
Mean change 3/03–3/31 = 0.46 μmol/L
Mean change 3/31–5/03 = 0.11 μmol/L
Mean change 1/29–5/03 = 0.79 μmol/L

shown in Table 4. Final adjustments to the ventilation system completed on March 9 probably contributed to the large decrease in the amount of airborne lead to which each person was exposed, as evidenced when comparing concentrations from the March 3 sampling with those from March 31 sampling. The data also show that the different booths on the range had very different air lead concentrations, both before and after changes were made to the ventilation system. The air lead concentrations in the classroom and control room dropped to below detectable levels after the first adjustment in the ventilation system.

The cumulative exposure to lead was directly related to the change in blood lead in Period 1 (Figure 1). We estimated lead exposure for the subjects not included in this figure by averaging the concentrations for the booths on either side of

TABLE 4—Air Lead Concentrations of Each Occupied Firing Range Booth, 8-hour Time-Weighted Averages in μg/m³, Colorado, 1987

	Date of Sample				
Booth	2/03 am	2/03 pm	3/03 am	3/03 pm	3/31
1	—	—	2917	—	2567
2	—	—	3095	1641	—
3	—	—	1823	994	811
4	2389	2688	3555	3398	837
5	PF	PF	2953	2359	850
6	639	513	1856	1883	553
7	2553	2426	4921	5589	1041
8	—	304	—	—	1006
9	—	—	4207	3269	1540
10	—	—	3706	4114	1875
Mean	1860	1483	3226	2906	1231
Classroom	175	170	<17	<17	<28
Control room	160	175	<17	<17	<23

Dual Image Shooting (March 31) = 205 μg/m³
— = Booth not being used.
PF = Pump failure
OSHA Standard is 50 μg/m³ for an 8-hour time-weighted average.

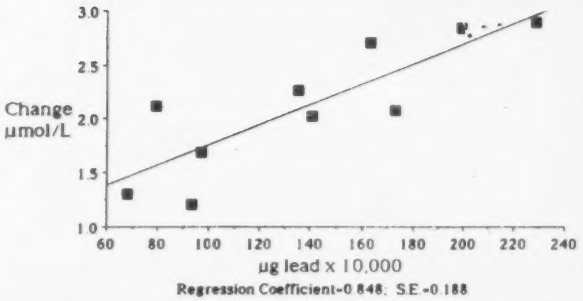

FIGURE 1—Change in Blood Lead Levels of Trainees* by the Cumulative Exposure to Lead in the Firing Range during Period 1 (January 29–March 3, 1987)
***Includes only the 10 trainees with complete data.**

the one with a failed pump, and redid this analysis; the results were very similar.**

Symptoms possibly related to lead exposure were reported by three trainees who complained of having a metallic taste in their mouth throughout most of the training.

The results of the air sampling using alternative ammunition (Table 5) indicated that substantial reductions in airborne lead were obtained through the use of the jacketed bullets. A 97 percent reduction occurred when using the copper-jacketed bullets.

Discussion

All of the blood lead levels collected on March 3, after four weeks of training, were above 1.45 μmol/L, considered elevated for non-occupational exposures; 14 were >1.93 μmol/L, the value at which the Occupational Safety and Health Administration (OSHA) requires active medical surveillance in occupational settings; and four were >2.90 μmol/L, the level at which OSHA requires medical removal from the source of exposure.[14] Even though most of the blood lead levels began to drop after March 3, only two trainees had a blood lead level below 1.45 μmol/L by the end of the study, and eight still had levels >1.93μmol/L.

The decrease in the blood lead levels after March 3 could be attributed to the lower mean time spent on the firing range and amount of ammunition fired by each recruit during March and April, final adjustments to the ventilation system, and five days of "dual-image" training where only one trainee at a time was on the range for a very short period of time and exposed to much lower levels of lead (205 μg/m³ versus an average of over 2,000 μg/m³—Table 4).

**Data available on request to authors.

TABLE 5—Air Lead Levels in μg/m³ Using Different Types of Ammunition, 8-hour Time-Weighted Averages, Colorado, 1987

Booth	Lead Bullets	Nylon-coated Bullets	Copper-jacketed Bullets
6	741.5	52.0	27.8
7	1800.9	82.3	43.7
8	1752.4	no data	30.5
9	1345.8	100.5	70.2
Mean	1410	78.3	43.1
Reduction	—	94.5%	97.0%

While the blood lead levels began to decrease after March 3, the FEP levels continued to increase. At the end of the study, 11 trainees had FEP levels >0.89 μmol/L, a level at which lead poisoning would be considered a more likely cause than iron deficiency anemia, which may also elevate the FEP level.[15–18] The FEP levels decreased for all nine trainees sampled about six weeks after completion of training. However, five still had FEP levels greater than 0.89 μmol/L.

The blood lead level reflects very recent and current lead absorption. It is not an indicator of the total body burden of lead. The FEP can reflect lead absorption over the previous three to four months and is indicative of biological effects on heme synthesis, thus being a better indicator of the total body lead burden.[15,19,20]

The overall mean air lead concentration for all booths before the March 31 sampling was >2,000 $\mu g/m^3$ (Table 4). At this concentration, the OSHA permissible exposure level of 50 $\mu g/m^3$ for an 8-hour time-weighted average[14] was exceeded within 12 minutes of shooting. The mean exposure level determined on March 31, after final adjustments to the ventilation system, was 1,231 $\mu g/m^3$. At this level, the permissible exposure level is exceeded in 20 minutes of shooting. The trainees were on the range a mean of 73 minutes per session in *Period 1*, 39 minutes per session in *Period 2*, and 41 minutes per session in *Period 3* (Table 2), all well above the time limit at which the OSHA permissible exposure level was exceeded given the range's air lead concentrations.

In recent years, the use of handguns has increased rapidly as a recreational activity, resulting in the formation of many gun clubs whose members use indoor firing ranges frequently.[7,8]*** This, in turn, has resulted in the increased construction and use of indoor firing ranges. The NRA and AFA estimate that there are about 800,000 competitive pistol shooters and 70 million privately owned handguns in the United States.*** Two major builders of indoor firing ranges—Detroit Armor and Caswell International—estimated that they have built about 50 public indoor firing ranges each year for the past 5–10 years, with the number increasing each year.*** In addition, the National Institute for Occupational Safety and Health estimates that there are 16,000 to 18,000 indoor firing ranges in the United States.

Interviews conducted with members of shooting clubs in the Denver area, as well as telephone communication with the NRA and the AFA, revealed that pistol club members commonly practice weekly as a group for as much as an hour or two at a time, each shooting approximately 100 to 200 rounds of ammunition each time. Serious competitive shooters reported practicing up to five hours per week shooting about 500 rounds of ammunition per week. If other indoor firing ranges have air lead concentrations similar to those found in this study then, based on our data, frequent users would be at risk for developing elevated blood lead levels and adverse health effects from the lead exposure.

At the very least, users should be made aware by means of conspicuous warning posters that exposure to lead from indoor firing ranges can cause adverse health effects. Improvements of ventilation systems may help to decrease air lead concentrations. However, such measures are extremely difficult and expensive to undertake and often do not alleviate the problem.[6] Another alternative is to limit the amount of exposure a person could receive by determining from air monitoring how long a person could remain on the range per day before the person's exposure exceeded the OSHA standard. While it is not clear whether a high exposure to lead in a short period of time is biologically equal to lower exposure over a longer period of time, this approach to controlling lead exposure has been suggested.[6] For some persons this may be acceptable, but for others, especially members of pistol teams, such a time limit may be unacceptable. Another option that also may not be popular with the public is requiring the use of respirators while using the indoor firing range. A final possibility, and one that may have the most potential, is to use jacketed ammunition, shown in this and other studies to be effective in reducing the air lead level in indoor firing ranges.[6,21]

***Personal communication with the National Rifle Association (NRA), the American Firearms Association (AFA), Detroit Armor, and Caswell International on October 5, 1987.

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Public Health Briefs

Accuracy of Women's Self-Report of Their Last Pap Smear

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Abstract: We compared interview data and physician records on when women last had a Pap smear in a sample of 98 rural Black women. We found 20 per cent of women could not accurately report on whether a Pap smear had been done within three years (sensitivity = 0.95, specificity = 0.47). Source of gynecological care and perceived barriers to obtaining a Pap smear but not education were associated with inaccurate reports. Self-report may be a misleading measure of Pap smear screening in comparable groups of women. (*Am J Public Health* 1989; 79:1036–1037).

Introduction

Women's own report of when they have had Pap smears is frequently the only source of information for clinicians who care for them, public health workers responsible for screening programs, and epidemiologists doing research.[1] However, some studies suggest that women may not recall accurately when they last received cervical cancer screening.[2–4] As part of a study examining patterns of Pap smear use among rural Black women, we assessed the accuracy of their recall of their last Pap smear and possible reasons for inaccurate self-report. We examined relations of source of care, educational level, and women's attitudes toward gynecological care to the accuracy of this recall.

Methods

We interviewed 149 Black women and reviewed records of their health providers in rural areas of three North Carolina counties. Approximately half the women were contacted by door-to-door canvasing of houses in low-income areas; the names of another third were provided by social workers in one county; the rest were friends and neighbors of the women referred by the social workers. During a structured, 30-minute home interview, two Black female interviewers asked the women about the date, place, and provider of their last Pap smear, their sources of medical care, attitudes toward gynecologic and preventive health practices, and perceived barriers to obtaining a Pap smear. After receiving informed consent from each woman, one of us (JAS) contacted providers by telephone to verify the women's reports on the date of their last Pap smear; a secretary or nurse provided the information after reviewing the woman's chart. We defined a "recent" Pap smear as one done within the past three years, based on the American Cancer Society's recommendation during the study period.[5]

We included in the questionnaire 17 questions about barriers to Pap smear use. We used a principal components factor analysis in a promax rotation to identify clusters of items (factors) that were related to each other.[6]

Results

Of the 149 women interviewed, 51 were excluded because they had had a hysterectomy (n = 28), had no source of medical care (n = 6), or were not listed on the rosters of any of the providers they named (n = 17). Our final sample of 98 women who had both a self-report and a provider report ranged in age from 16 to 75 years (mean 35), and in education from less than six to 16 years (mean 10.8; 22 per cent ≤9 years, 21 per cent 10 or 11 years, 57 per cent ≥12 years); their mean family income was $11,040 per year. Their demographic characteristics did not differ appreciably from 1980 census data for the three counties.

Women's self-reports correlated poorly with providers' reports on date of their last Pap smear (Table 1). Although there were errors in both directions, women usually overestimated how recently they had had a Pap smear. Using whether a Pap smear had been done in the past three years as a cutpoint, there was a 20 per cent error rate (Table 2). Women's self-report had a sensitivity of 95 per cent and a specificity of 47 per cent. Agreement between patient and physician was only fair (Kappa = 0.46).

As shown in Table 3, women seeing nurse practitioners were more likely to report accurately when their last Pap smear had been done than women seeing internists or family physicians. There were no clear differences in accuracy of reporting by age, income, education, or method of recruitment into the study.

Factor analysis identified two factors which we used to construct two variables involving six of the 17 potential barriers (the other 11 items showed no further discriminating

TABLE 1—Agreement between Provider and Women's Report of Number of Years since the Last Pap Smear

Women's Report	Provider's Report					
	<1	1–2	3–6	≥7	Never	
<1	42	6	5	1	6	60
1–2	3	12	2		3	20
3–6	1	1	5	1		8
≥7	1			5		6
Never			1		3	4
	47	19	13	7	12	98

Pearson Correlation Coefficient = 0.496

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TABLE 2—Agreement between Provider and Women's Report of Pap Smear within Three Years

Women's Self-Report	Provider's Report		
	Recent Pap	No Recent Pap	Total
Recent Pap	63	17	80
No Recent Pap	3	15	18
Total	66	32	98

Self-Report: Sensitivity = 95%
Specificity = 47%
Positive predictive value = 79%
Negative predictive value = 83%

TABLE 3—Percentage of Women Reporting Accurately by Type of Gynecologic Provider and Perceived Barriers

	N	Accurate Self-Reports %	OR	95% CI
Overall		80		
Gynecologic Provider				
Obstetrician Gynecologist	15	87	1.00	*
Nurse Practitioner	27	93	1.90	0.24–15
Internist Family Practice	41	68	0.33	0.06–1.68
Other	15	80	0.62	0.09–4.4
Pelvic examination beliefs				
Important**	30	67	0.34	0.13–0.91
Not important	68	85	1.00	*
Perceived trouble of examinations				
Important**	35	69	0.36	0.19–0.71
Not important	63	86	1.00	*
Recruitment method				
Door to door	46	76	1.00	*
Other	52	83	1.50	0.77–2.9

*Baseline
**Important = agreeing that at least two-thirds of the component items are important barriers to getting Pap smears

ability). The "pelvic examination barriers" factor contained three items: not liking to be examined by a male; pelvic examinations being unpleasant; pelvic examinations being embarrassing. The 30 women who felt that at least two of three of these barriers were important reasons why women did not get regular Pap smears gave less accurate reports than other women did. The factor labeled "trouble getting examinations" contained three items: seeing a doctor takes a lot of time; not wanting "to look for trouble"; meaning to get a Pap smear but "just not getting around to it". The 35 women who felt that at least two of three of these barriers were important reasons why women did not get regular Pap smears gave less accurate reports of when their last Pap smear had been done than did other women.

Discussion

A limitation of these data is that providers' records are not necessarily the "gold standard" for when the last Pap smear was done. It is possible that the medical record contained inaccurate data, or that dates were overlooked in the record review. If many Pap smear reports were overlooked, our estimate of the specificity of self-report would increase.

Women seen by internists and family physicians were least likely to report accurately about the recency of their last Pap smear. It may be that more of these women had pelvic examinations done for reasons other than Pap smears and either confused the term "pelvic examination" with "Pap smear" or assumed that a Pap was done routinely during every pelvic examination. Such confusion is important if women believe that they have had a Pap smear when they have not and may well prevent them from receiving one when they need it.

More than confusion, however, is involved; those women with inaccurate reports were also more likely to perceive significant logistical barriers to getting a Pap smear and to view pelvic examinations as embarrassing or unpleasant. Perhaps women with negative perceptions remember their last Pap smear more vividly. Vivid experiences are recollected as more recent, regardless of when they actually occurred and this may account for some of the inaccuracy of self-report.[7]

Somewhat surprisingly, a woman's educational level was not related to whether or not she recalled accurately when her last Pap smear had been done. Others also have found education to be unrelated to the accuracy of survey data.[4,8]

Although one may not be able to generalize these results, they highlight the danger of relying exclusively on self-reports for determining when a woman last received a Pap smear.

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Nurse-Midwifery Practice in the United States, 1982 and 1987

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Abstract: Surveys of certified nurse-midwives living in the United States were conducted on behalf of the American College of Nurse-Midwives for 1982 and 1987. Most of the respondents were providing prenatal and family planning care. The 1982 respondents conducted 1.8 percent of the deliveries that took place in the United States during 1982 and 1988 respondents conducted 2.5 percent of all 1987 United States deliveries. (*Am J Public Health* 1989; 79:1038–1039.)

Introduction

A certified nurse-midwife (CNM) is a registered nurse with advanced training in midwifery who possesses evidence of certification by the American College of Nurse-Midwives (ACNM). Nurse-midwifery practice encompasses responsibility for management of health care for women and newborns who do not have medical problems. Care for women is provided during the antepartal, intrapartal, and postpartal as well as the preconceptional and postconceptional phases of the reproductive cycle. Nurse-midwifery care occurs within a health care system that provides for medical consultation, collaborative management, and medical referral.[1,2]

United States resident CNMs were surveyed at the end of 1982 and in January 1988, to determine what services were being provided by these CNMs.

Methods

Survey questionnaires were mailed to 2,086 CNMs known to be living in the United States in 1982, 81.8 percent of the 2,550 ever-certified by the ACNM as of July 1, 1982, and to 3,846 CNMs known to be living in the United States in 1988, 97.1 percent of the 3,959 ever-certified by the ACNM as of October 1, 1987. An average of 360 registered nurses graduated from nurse-midwifery education programs accredited by the ACNM each year between 1982 and 1987. These graduates account for the additional CNMs surveyed in 1988. Participants were asked to respond to conditions and events pertaining specifically to 1982 and 1987, respectively. Identical questions about services provided were contained in both questionnaires. Usable questionnaires were returned by 1,584 (75.9 percent) of the CNMs in 1982 and 2,178 (56.6 percent) of the CNMs in 1988.

Results

A total of 1,065 (67.2 percent) of the 1982 and 1,527 (70.1 percent) of the 1988 respondents were practicing nurse-midwifery. Prenatal and family planning care were the services provided by the largest numbers of respondents (Table 1). Major increases in the proportions of respondents who conducted deliveries in hospital birth centers and who conducted multiple gestation deliveries were noted between 1982 and 1987.

The 937 respondents conducting deliveries in 1982 reported conducting 66,881 deliveries during that year, or about 1.8 percent of the 3,704,000 deliveries in the United States.[3] The 1,261 respondents who reported conducting deliveries in 1987 reported conducting 94,031 deliveries during that year, or about 2.5 percent of the 3,829,000 deliveries in the United States.[4] These CNM-conducted deliveries include deliveries conducted by nurse-midwifery students under the direction of CNMs.

The actual number of deliveries conducted by CNMs in the United States is unknown. The National Center for Health Statistics record numbers of deliveries conducted by midwives from information on birth certificates; however, deliveries include those conducted by lay midwives as well as CNMs. At least 10 percent of the 1982 and 1987 CNM respondents indicated on the questionnaire that they do not sign birth certificates for the babies they deliver.

Most deliveries conducted by respondents took place in hospitals. Only 12.5 percent (133) of the 1982 respondents and 10.7 percent (164) of the 1988 group conducted deliveries in homes; 12.6 percent (134) of the 1982 respondents and 11.1 percent (170) of the 1988 group conducted deliveries in out-of-hospital birth centers. Although only 64.8 percent (989) of the 1988 respondents conducted deliveries in traditional hospital settings compared with 71.4 percent (760) of the 1982 respondents, the percent who conducted deliveries in hospital birth centers increased from 30.0 (319) to 35.4 (541).

Participation in the care of some high-risk women during labor and delivery is suggested by the number of 1988 respondents who applied internal uterine pressure monitoring devices (1,063) or fetal scalp electrodes (988), rotated the fetus manually during labor (421), began oxytocin augmentations without physician consultation (122), interpreted electronic fetal monitoring strips (223), conducted breech delivery (141) or multiple gestation delivery (301), or removed the placenta manually when indicated (800). Some respondents also indicated that they repaired fourth degree lacerations or epitiotomy extensions (408), and cervical lacerations (552), as well as performed newborn intubation by laryngoscope (440) when indicated.*

For the purposes of both surveys, continuity of care was defined as care provided by an individual CNM, either alone or with CNM or physician colleagues, to a defined caseload of women through the antepartal, intrapartal, and postpartal phases of pregnancy. By 1987 at least 80 percent of the respondents working in all types of organizations reported providing continuity of care and at least 70 percent of the women they cared for received this care. Almost all of the respondents who worked in private practices provided continuity of care.

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*Detailed information about the activities performed by CNMs can be obtained from the American College of Nurse-Midwives, 1522 K Street, NW, Suite 1120, Washington, DC 20005.

TABLE 1—1982 and 1987 ACNM Survey Participants Practicing Nurse-Midwifery by General Area of Patient Care Management and Specific Activities Performed within the General Area of Patient Care Management

General Area of Patient Care Management	Specific Activity	Percent Who Performed Activity 1982 (N:1,065)	Percent Who Performed Activity 1987 (N:1,527)
Prenatal	Provide prenatal care	96.4	89.0
	Diagnose pregnancy	91.0	81.0
Labor	Provide care during labor	89.1	79.4
Delivery	Conduct deliveries	88.0	82.6
	Conduct deliveries in private homes	12.5	10.7
	Conduct deliveries in a hospital birth center	30.0	35.4
	Conduct deliveries in a non-hospital birth center	12.6	11.1
	Conduct deliveries in a traditional hospital setting	71.4	64.8
	Conduct breech deliveries	12.6	9.2
	Conduct multiple gestation deliveries	12.6	19.7
	Remove the placenta manually when indicated	55.3	52.4
	Repair third degree lacerations and episiotomy extensions	69.3	60.8
	Repair fourth degree lacerations and episiotomy extensions	36.3	26.7
	Sign birth certificates for nurse-midwifery conducted deliveries	76.9	73.8
Immediate Newborn	Provide care for the newborn at birth	87.3	54.2
	Perform intubation by laryngoscope when indicated	32.0	28.8
	Perform routine physical examinations	49.6	30.4
Postpartum	Provide postpartum care	83.5	80.3
	Discharge postpartum women without physician consultation	73.4	61.8
Family Planning	Provide family planning care	90.6	83.6
	Prescribe and supervise the use of oral contraceptives	85.5	79.2
	Insert IUDs and provide follow-up care	69.5	34.6
Well Woman Gynecology	Perform gynecologic histories and physical examinations	77.7	72.2

TABLE 2—Percent of 1982 and 1987 ACNM Survey Participants Who Conducted Delivery by Type of Work Organization and Percent of CNM-Conducted Deliveries

Type of Work Organization*	1982** Percent of CNMs (N:890)	1982** Percent of CNMs Conducting Deliveries (N:66,881)	1987*** Percent of CNMs (N:1,261)	1987*** Percent of CNMs Conducting Deliveries (N:94,031)
Hospital or Medical Center	32.2	30.9	63.4	70.7
Public Health Agency	5.2	4.2	18.5	20.4
US Military or Public Health Service	5.6	8.2	6.4	9.1
Prepaid Health Care Plan	4.8	5.1	13.1	16.2
University Health Service	1.2	1.0	3.6	2.9
Private Practice with One or More Physicians	17.4	21.5	25.4	25.6
Private Nurse-Midwifery Practice	21.9	19.1	19.9	13.8
Other	11.6	10.1	8.7	7.6

*Unknowns excluded.
**1982: primary work site only.
***1987: all who worked in type of organization whether primary, secondary, or tertiary work site.

Discussion

Nurse-midwifery continues to grow as a health profession in the United States. The number of nurse-midwives ever certified by the ACNM increased by 55.3 percent between 1982 and 1987. Although a lower percent of 1988 respondents conducted deliveries in traditional hospital settings, a much higher percent conducted deliveries in hospital birth centers. This is a reflection of more readily available options for birth sites within hospitals offered to women at low-risk of delivering obstetrical complications. About the same proportion of 1982 and 1988 respondents conducted home deliveries or deliveries in out-of-hospital birth centers.

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Left-Handedness and Accident-Related Injury Risk

STANLEY COREN, PHD

Abstract: Self-reported injuries among left-handed and right-handed people were compared in a survey of 1,896 college students in British Columbia, Canada. Left-handers were more likely to report having an injury requiring medical attention during the last two years (OR = 1.89, 95% CI = 1.39, 2.58). Relative risk was highest for left-handed males when driving motor vehicles (OR = 2.35, CI = 1.25, 4.43). Regardless of handedness, males had slightly higher relative risks of injury than females. (*Am J Public Health* 1989; 79: 1040–1041.)

Introduction

Several reports show that the percentage of left-handers in the population decreases with increasing age,[1–4] with about 13 percent left-handedness in samples of 20 year olds, diminishing to 5 percent for those in their 50s, and to virtually 0 percent for samples age 80 and older.[1] Evidence suggests that this change in the frequency of left-handers is not explainable by historical trends in social pressure for right hand use.[1–3] In the belief that the absence of left-handers from the oldest age groups reflected higher risk of mortality for sinistrals, Halpern and Coren[5] conducted a study that suggested an association between left-handedness and shorter life-span. Using archival records from 2,271 deceased professional baseball players (excluding those with evidence for switched or inconsistent handedness), they found that the annual survival rate of right-handers exceeded that of left-handers by an average of about 2 percent. They suggested three possible causes:

- Left-handedness is sometimes associated with prenatal and perinatal stressors (e.g., low birthweight, prolonged labor, birth-related anoxias, Rh incompatibilities, and older mothers[1,6,7]) which may have sequelae that result in reduced survival ability.
- There are demonstrable associations between left-handedness, high levels of prenatal hormones, allergies, and immune disorders that may increase the risk of early mortality.[8,9]
- An environmental factor, the technological and behavioral bias toward dextrality, may be involved. Porac and Coren[1] collected numerous anecdotal reports that, because tools, machinery and even traffic patterns have been designed for the convenience of right-handers, left-handers may be more subject to accidental injuries which, individually or cumulatively, may result in reduced longevity.

The study reported below investigates the effects of handedness on injury risk.

Methods

A retrospective self-report procedure was employed. A sample of 1,896 students from the University of British Columbia (1,251 Caucasian, 567 Oriental, and 78 Black, American Indian or other, mean age 21.0, S.D. 5.62) were assessed in intact classroom units. All were tested for handedness using a four-item self-report inventory that assessed the hand used to perform common activities, specifically: drawing, throwing a ball, dealing cards and using an eraser on paper. This set of items has been shown to have a 97 percent concordance with behavioral testing and to be reliable over test-retest periods of a year.[1,10,11] In addition, individuals were asked to report if, in the previous two years, they had had any accidents in five categories of activity: when using tools or implements, while driving in traffic, when at home, when at work, or while participating in sports. Only accidents resulting in injuries that required medical attention were considered. No data were collected on medical events other than injuries, and medical confirmation of reported injuries was not sought.

Results

Analysis of the handedness patterns revealed 1,718 (90.5 percent) right-handers and 180 (9.5 percent) left-handers in this sample, which is similar to the usually obtained handedness distribution for this age group.[1]

Table 1 presents the distribution of injuries as a function of handedness as percentages and frequencies, while Table 2 presents estimates of relative risk in the form of odds ratios and their 95 percent confidence intervals. The relative risk of all injuries were elevated for left-handers. Although males tended to report more accident-related injuries than females overall (OR = 1.81, CI = 1.50, 2.18), left-handers had higher relative risks for both genders.

Among individuals reporting injuries in more than one category of activity, male and female left-handers were at higher risk.

Consideration of specific injury categories must be done with some caution, since we are partitioning the data into somewhat sparse groupings, and the relative distributions will be highly asymmetrical given the fact that the vast majority of the sample are right-handed and the bulk of the sample have not suffered from an injury in any given category. Table 2 shows higher relative risk of injury for left-handers in each of the five categories, although the only individual activity classification for which the elevation of risk is statistically significant is for injuries while driving a vehicle. The binomial test for all five comparisons in the same direction has $p = 0.04$. The same pattern appears with data segregated according to gender.

Discussion

These data suggest that overall (ignoring specific categories of activity) left-handers have increased risk of accident-related injuries, and higher risk of being injured in multiple categories of activity. Alternatively, the patterns seen could reflect a greater likelihood of left-handers to remember or report accidents, but we know of no data to suggest such a bias.

The most likely reason for the increased risk of injuries in left-handers comes from implicit and explicit biases of the environment toward maximal convenience of the right-

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TABLE 1—Percentage of Individuals Suffering an Accident-Related Injury Requiring Medical Attention, as a Function of Handedness (frequencies in parentheses)

	All Cases (N = 1896)		Females (N = 1086)		Males (N = 810)	
Accidential Injury Category	Left-handed (180)	Right-handed (1716)	Left-handed (96)	Right-handed (990)	Left-handed (84)	Right-handed (726)
At least one injury any category	51.7% (93)	36.1 (619)	42.73 (41)	30.6 (303)	63.0 (53)	43.6 (316)
More than one injury category	20.0 (36)	12.3 (211)	15.5 (15)	9.6 (95)	25.0 (21)	16.0 (116)
Work-Related	15.4 (28)	12.8 (220)	11.6 (11)	10.5 (104)	20.0 (17)	15.9 (115)
In the Home	12.6 (23)	9.0 (154)	10.6 (10)	9.0 (89)	15.5 (13)	9.0 (65)
Sport-Related	31.6 (57)	27.9 (479)	25.5 (25)	21.6 (214)	39.2 (33)	36.4 (264)
Driving a Vehicle	10.3 (19)	6.0 (103)	5.3 (5)	4.7 (46)	16.6 (14)	7.8 (57)
Using tools or implements	7.5 (14)	5.2 (89)	6.4 (6)	3.8 (38)	8.3 (7)	7.0 (51)

TABLE 2—Odds Ratios* for Accident-Related Injuries in Left-Handers with 95% Confidence Intervals

	All Cases		Female		Male	
	OR	CI	OR	CI	OR	CI
At least one injury any category	1.89	1.39–2.58	1.69	1.10–2.59	2.22	1.39–3.54
More than one injury category	1.78	1.20–2.64	1.75	0.97–3.15	1.75	1.03–2.99
Work-Related	1.25	0.81–1.92	1.10	0.57–2.13	1.35	0.76–2.38
In the Home	1.49	0.93–2.37	1.18	0.59–2.35	1.86	0.98–3.55
Sport-Related	1.20	0.86–1.67	1.28	0.79–2.06	1.13	0.71–1.80
Driving a Vehicle	1.85	1.10–3.10	1.13	0.44–2.91	2.35	1.25–4.43
Using tools or implements	1.54	0.86–2.77	1.67	0.69–4.06	1.20	0.53–2.74

*Defined as (injured LH/uninjured LH)/(injured RH/uninjured RH)

handed majority. Everyday implements, such as scissors, gearshifts, and can openers, even the direction in which the threading of screws is angled, are biased toward right-handed use. Traffic patterns are designed to utilize the clock-wise turning bias of the right-hander. Many power tools, in their usual configurations, provide for free use of the right hand to manipulate materials, while restricting movements and utility of the left-hand, as can be seen in lathes, band saws, and some milling machines. Thus, to function in the right-handed world, the left-hander must either work with his non-dominant and less proficient right hand, or must adopt body postures and manipulation patterns which are at variance with the design of the machines. Such activities place the sinistral at a higher risk of suffering an accidental injury. This increased risk may be a significant factor in determining the longevity of sinistrals as a group (since injuries have been shown to be the underlying cause of death for more than 95 percent of individuals aged less than 55 years[11]) and may account for the relative rarity of left-handers in older age groups.

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Deaths, Injuries, and Evacuations from Acute Hazardous Materials Releases

SUE BINDER, MD

Abstract: We examined reports from three surveillance systems of 587 acute releases of hazardous materials in 1986. These releases resulted in at least 115 deaths, 2,254 injuries, and 111 evacuations. Only eight (1 percent) of the 587 events were common to all three systems. Estimates of the public health consequences of hazardous materials releases could be improved by enforcing existing laws, modifying report forms, and validating collected information. (*Am J Public Health* 1989; 79:1042–1044.)

Introduction

The Superfund Amendments and Reauthorization Act of 1986 contains extensive provisions about emergency planning and the community's right-to-know about toxic chemical releases. A 1987 survey showed that information about public health consequences of such releases was generally not available at the state and regional levels (Sue Binder, unpublished data). Accordingly, we evaluated the three largest national sources of data that record deaths and injuries from acute chemical releases—the National Response Center (NRC), the Department of Transportation (DOT) Hazardous Materials Information System (HMIS), and the Acute Hazardous Events (AHE) data base. Our goal was to determine if these data bases could yield useful public health information. Characteristics of these data sources are discussed in the Appendix.

Methods

This assessment focused on incidents occurring in 1986, the last year for which complete data were available. All 1986 events coded in AHE and HMIS as having resulted in death, injury, or evacuation were obtained. Since NRC did not code evacuations in 1986, only events resulting in death or injury from hazardous materials releases were included. Events were coded as transportation-related if they appear in HMIS, were stated explicitly in NRC to involve a vehicle carrying a toxic chemical or to involve loading or unloading of a vehicle, or were coded in AHE either as occurring in transit or as occurring during loading or unloading and as involving a vehicle not in transit.

The data were converted to a Statistical Analysis System (SAS) format.[1] Events appearing in more than one system were identified and collapsed into one report, and then all reports were merged to form the Public Health Consequences data base. Descriptive analyses were conducted using SAS procedures.[1]

Results

The Public Health Consequences data base included 587 releases resulting in deaths, injuries, or evacuations (Figure 1). Only eight (1 percent) events appeared in all three data bases. These eight events all involved transportation (by definition, since they appear in HMIS) but were otherwise unclustered by type of chemical or in the numbers of people injured or killed.

Fifty-eight events resulted in a total of 115 reported deaths. Thirty-six of these events resulted in a single death each; the largest number of deaths in a single event was 14. Four hundred ninety-six events resulted in a total of 2,254 injuries. Three events resulted in over 100 injured persons each. Although 111 evacuations were identified from HMIS and AHE, the numbers of persons evacuated were missing for 75 percent of evacuations.

The chemicals involved in these incidents were unknown in 23 (4 percent) of the 587 events. For many events, only the class of chemical involved (e.g., corrosive liquid) was specified. Table 1 lists all chemical entries appearing more than five times in a data base.

Three hundred twenty-eight events, 56 percent of the events in the Public Health Consequences data base, involved an in-transit vehicle or an activity related to transportation (e.g., loading or unloading a vehicle). One hundred nine events represented crashes, derailments, and vehicular overturns—events that might occur in areas unprepared to deal with hazardous materials spills.

Discussion

The merged reports from the three systems indicate that there are an average of 1.6 hazardous materials incidents a day in the United States that result in either death, injury, or evacuation; almost once every three days, there is a release from a crash, derailment, or overturned vehicle that results in a similar outcome.

According to the Public Health Consequences data base, the number of people killed from hazardous materials releases in 1986 was 115 and the number injured was 2,254. Although these estimates are probably the most accurate so far, they represent the minimum number of persons killed and injured. A major reason is the mandated limits in scope of the

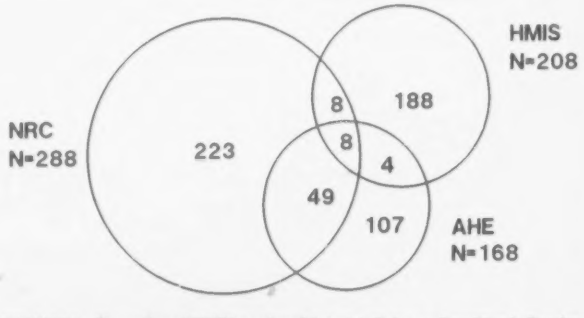

FIGURE 1—Reporting of 587 Hazardous Materials Releases Resulting in Deaths, Injuries, or Evacuations by Data Base, USA, 1986
NRC = National Response Center Data Base
HMIS = Hazardous Materials Information System
AHE = Acute Hazardous Events Data Base

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TABLE 1—Chemicals and Chemical Groups Associated with Releases Resulting in Deaths, Injuries, or Evacuations, by Data Base, 1986*

Public Health Consequence Data Base	National Response Center Data Base	Hazardous Materials Information System	Acute Hazardous Events Data Base
1. Natural gas (88)	Natural gas (88)	Sulfuric acid (21)	Chlorine (35)
2. Chlorine (38)	Chlorine (18)	Gasoline (16)	Ammonia (16)
3. Gasoline (33)	Gasoline (17)	Hydrochloric acid (14)	Hydrochloric acid (11)
4. Sulfuric acid (30)	Diesel oil (12)	Sodium hydroxide (8)	Sulfuric acid (11)
5. Hydrochloric acid (25)	Ammonia (6)	Corrosive liquids (8)	
6. Ammonia (21)	Sulfuric acid (6)	Cleaning liquids (6)	
7. Sodium hydroxide (13)	Miscellaneous oils (5)	Phosphoric acid (6)	
8. Diesel oil (12)		Poisonous liquids (6)	
9. Corrosive liquids (8)		Ammonia (5)	
10. Nitric acid (7)			
11. Phosphoric acid (7)			
12. Cleaning liquids (6)			
13. Flammable liquids (6)			
14. Formaldehyde (6)			
15. Poisonous liquids (6)			
16. Sulfur dioxide (6)			
17. Toluene diisocyanate (6)			
18. Miscellaneous oils (5)			

*Number of appearances is in parentheses. Only those appearing in the data bases five or more times are listed.

contributing systems. For example, HMIS excludes incidents involving intrastate carriers. A second problem is that the data bases probably do not capture all events falling within their mandate.[2,3] The fact that only eight of 595 events (1 percent) appear in all three data bases probably results both from the different areas of emphasis and failure to report by responsible parties.

In addition to incompleteness, the three component data bases have other limitations for evaluating public health impacts. The accuracy of the death and injury information is unknown, particularly for NRC and AHE. Reporting biases may result because certain industries or companies are more likely to report than others. Responsible parties may wish to present themselves in the best possible light, and may therefore tailor their reports accordingly.

The described characteristics of events in the data bases should be interpreted with caution. In particular, the results in Table 1 should not be interpreted as indicating the frequency or even the relative frequency with which a chemical was associated with deaths, injuries, or evacuations, since compliance with reporting requirements and accuracy of identification varies with different chemicals. The chemicals listed in Table 1 also reflect differences in emphasis among reporting systems. Natural gas releases, many from pipeline ruptures or occurring within private residences, were common in NRC. The lack of events involving natural gas, gasoline, and oils in AHE reflects its exclusion of chemicals not covered in Superfund legislation.

Although collecting data on hazardous materials spills is difficult, costly, and time-consuming, it could be improved by enforcing laws on incident reporting; obtaining more specific information about deaths, injuries, and evacuations; and validating collected information. Accurate and complete data on deaths, injuries, and evacuations resulting from acute hazardous materials releases could be used in many ways. Linking these data with information about commodity flows and geographic distributions of chemicals would allow calculation of rates to identify industries, types of facilities, activities, and areas of the country with higher risks of having releases with public health consequences. These rates would make it possible to generate hypotheses about why these events are occurring and to intervene appropriately. The identification of high-risk chemicals, facilities, and activities would also be of use to persons planning emergency response activities and to concerned communities. Although the available data are useful for providing minimum estimates of some of the public health consequences of unintentional hazardous chemical releases, improved surveillance would be an important step toward understanding the problem and, ultimately, preventing it.

ACKNOWLEDGMENTS

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APPENDIX

Description of Data Sources

The National Response Center (NRC), staffed and administered by the US Coast Guard on behalf of 15 participating federal agencies, has been collecting information on chemical releases and serving as a coordinating center for emergency response since 1974. The party responsible for the release is required to notify NRC about hazardous chemical and radiological releases exceeding certain quantities; hazardous materials releases from pipeline failures or cargo in transit resulting in specified outcomes, including death or serious injuries; and certain releases of toxic, corrosive or flammable gas, liquefied natural gas, or gas from a liquefied natural gas facility. By law, NRC must be notified by telephone within 24 hours for transportation-related events or as soon as practical for other events. Since this system is based on initial reports, the information is often incomplete. Much of the information is entered directly into a computer as long, uncoded fields of text. The data do not always indicate whether injuries or deaths resulted from the released chemical or from impact, as from a crash. Until recently, evacuations were uncoded. Data are generally neither verified nor updated.

The Hazardous Materials Information System (HMIS)—the main Department of Transportation (DOT) surveillance system for chemical releases since 1971—contains information about releases of hazardous materials in transit and during transportation-related loading, unloading, or temporary storage. By law, carriers must file written reports to HMIS within two weeks of a chemical release incident. These reports include information about the type of transport,

cause of the incident, container structure and failure, and deaths and injuries resulting from the cargo. DOT attempts to validate death and injury data. Specifically excluded from reporting requirements are releases of small quantities of certain consumer commodities, and releases from motor carrier firms doing solely intrastate business and from certain water transporters. Automobiles striking storage tanks and certain transportation-related spills at fixed facilities are also excluded.

The Acute Hazardous Events (AHE) data base was begun in 1985 and uses the NRC as its main source of data. However, data are also included from selected state governments, the Environmental Protection Agency (EPA) Region 7, some newspapers and wire services, and other sources. Information collected includes cause of event, activity taking place during the event, and type of property damaged. Attempts are made to eliminate deaths and injuries not caused by hazardous materials. Because emphasis was placed upon events involving chemicals covered by Superfund legislation and air releases from fixed sites, many events which appear in the NRC data base are excluded. AHE is maintained and augmented by EPA and its contractors, primarily Industrial Economics, Incorporated, and has been updated through 1986.[4] If events appear in more than one source, they are checked for consistency; otherwise, data are not validated.

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Formaldehyde Exposures from Tobacco Smoke: A Review

THAD GODISH, PHD

Abstract: Reports of formaldehyde levels in mainstream, sidestream, and environmental tobacco smoke from nine studies are reviewed. Considerable disparity exists between formaldehyde production rates determined from mainstream-sidestream studies and those reporting levels in environmental tobacco smoke. Tobacco smoke does not appear to increase vapor-phase formaldehyde levels significantly in indoor environments, but formaldehyde exposure in mainstream smoke may pose a risk of upper respiratory system cancer and increase the risk of cancer in smokers. (*Am J Public Health* 1989; 79: 1044–1045.)

Introduction

Formaldehyde is a major oxidation by-product of combustion processes including tobacco smoking. It is produced in both the mainstream (MS), and sidestream smoke (SS), and has been reported in substantial levels in environmental tobacco smoke (ETS).

Formaldehyde levels in mainstream, sidestream, and environmental tobacco smoke reported by a number of investigators are summarized in Table 1. Reported studies vary in testing methodologies and expression of concentrations. Concentration units are those originally reported and those calculated and standardized by the author from original data, assuming a smoking rate of 35 ml/puff and 10 puffs/cigarette.

As seen in Table 1, formaldehyde concentrations in mainstream smoke[1–4] ranged from about 10 μg/cigarette to over 100 μg/cigarette. Differences in concentrations reflect differences in tobacco type and brand. Higher average concentrations reported by the Surgeon General in 1986[3] reflect those of regular non-filter cigarettes.

Sidestream vapor-phase formaldehyde concentrations also varied somewhat. Ayer and Yeager[5] reported 15–48 ppm. Hoffman's observations ranged from nondetectable to 34.2 μg/cigarette, with an average of 12.1 μg/cigarette for 16 different brands.[4]

Room or large chamber formaldehyde levels associated with environmental tobacco smoke[6–9] indicate that formaldehyde concentrations in such rooms are high. For example, in the studies of Howlett, *et al*,[8] one cigarette smoked in an environmental chamber caused the formaldehyde level to increase to 0.21 ppm within a half hour. Formaldehyde production rates calculated from ETS concentrations (Table 1) are substantially higher (one to two orders of magnitude) than those reported for MS, SS, and MS-SS combined.

The considerable disparity in formaldehyde production rates determined from MS-SS and ETS studies suggests differences due to methodologies employed in sampling and analysis. In the mainstream-sidestream smoke studies reported by the Surgeon General[2,3] and by Hoffman,[4] gas and particulate phase materials were separated by high-efficiency filtration. In studies by Weber, *et al*,[6] no attempt was made to remove particulate phase materials. Sundin[7] employed particulate phase filtration of unknown efficiency. Attempts to remove particulate phase materials in ETS samples were not reported by Howlett, *et al*,[8] and Klus, *et al*.[9]

In mainstream-sidestream smoke studies,[2–4] smoke samples were analyzed by the 2,4 dinitrophenylhydrazene-HPLC method which is specific for free formaldehyde. The chromatropic acid method[10] on the other hand was used in the studies of Weber, *et al*,[6] and Sundin[7]; it is likely to have been employed in the two other environmental tobacco smoke studies as well because it is the dominant method used to determine formaldehyde concentrations in air. In the chromatropic acid method, formaldehyde forms a stable addition product on sample collection in sodium bisulfite solution. On analysis, the addition product is destroyed yielding free formaldehyde. Any solution which contains free formaldehyde, a formaldehyde addition product, or organic compounds which produce formaldehyde on sulfuric acid destruction will test positive for formaldehyde.

On analysis with the chromatropic acid method, the particulate phase of tobacco smoke has been shown to contain appreciable quantities of formaldehyde.[5] This formaldehyde may be present as free formaldehyde dissolved in liquid water or it may be produced by the destruction of formaldehyde addition products and possibly other organic

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Elimination of Cotinine from Body Fluids: Disposition in Smokers and Nonsmokers

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Abstract: We have evaluated differences in the elimination of cotinine, a major nicotine metabolite, in smokers who quit smoking and never-smokers who were exposed to environmental tobacco smoke (ETS) under controlled conditions. The mean biological half-life of cotinine in urine, collected from the nine smokers was 16.5 ± 1.2 h, in never-smokers exposed to ETS, 27.3 ± 1.9 h. Differences in the mode of uptake and absorption of nicotine and possible differences in nicotine metabolism may play roles in the clearance rate differences between smokers and nonsmokers. (*Am J Public Health* 1989; 79:1046–1048.)

Introduction

The elimination of nicotine, by smokers, users of other tobacco products, and by nonsmokers exposed to environmental tobacco smoke (ETS), is a variable metabolic process.[1–3] In a previous study, we observed that cotinine elimination in nonsmokers exposed to ETS tended to take longer than that seen in active smokers who quit smoking.[4] The slower clearance in nonsmokers of a terminal nicotine metabolite suggested that other tobacco smoke constituents or their metabolites might also remain in the bodies of nonsmokers for longer time periods.

Differences in the induction of hepatic and possibly pulmonary enzymes can play a role in the diverse metabolic rates observed between individuals.[5–11] Additionally, variations in the site and rate of absorption of nicotine can affect first pass metabolism by the liver and influence nicotine conversion as well as nicotine clearance.[8–10] Therefore we designed the current study to evaluate possible differences in the rate of cotinine elimination in smokers and ETS-exposed never-smokers.

Methods

Nine smokers between 23 and 38 years of age abstained from cigarette smoking for the duration of the study. Their prior smoking behavior ranged from self-reported use of six to 30 cigarettes per day. Smoking status was confirmed by urine cotinine concentrations. All urine voids were collected in four daily aliquots for five days beginning with cessation of smoking.

Ten nonsmokers who reported that they were never-smokers (26–45 years of age) were exposed to ETS for 80 minutes twice daily in two successive days in a specially constructed chamber. Five subjects were exposed at each time to the sidestream smoke produced by four Kentucky 1R1 Reference cigarettes which were continuously smoked throughout the exposure period. The chamber conditions have been described previously.[12] Briefly summarized, the chamber is 16.3 m^3 with six air changes per hour which corresponds to the average ventilation condition of offices in the United States. The pollution levels of selected smoke constituents were approximately 25 ppm CO, 0.9 ppm NO_x, and 160 $\mu g/m^3$ formaldehyde. These are considered high pollution levels. Urine samples were collected before exposure and in four daily aliquots as described for the smokers following cessation.

Each sample was analyzed for cotinine by radioimmunoassay (RIA) using specific antisera produced in rabbits.[13,14] Creatinine in urine was determined on a Kodak Ektachem 400 Clinical Chemistry Analyzer.[15] Creatinine excretion was used to normalize urine cotinine values.[14]

The volumes of smoke exhaled during smoking or ETS exposure are variable as well as the efficiency of lung clearance. These factors and the amount of matter expectorated or swallowed and absorbed from the digestive tract will affect bioavailability. In light of these variables, calculations of the dose of smoke taken up by subjects would be inappropriate. We have therefore assessed exposure by terminal cotinine half-lives ($t_{1/2}\beta$) estimated by linear regression of the log concentrations against time using an interactive computer program designed for the analysis of drug pharmacokinetics.[16]

Results

Data on the elimination of cotinine in the urine of smokers and ETS-exposed nonsmokers are presented in Table 1. The mean of the $t_{1/2}\beta$ for cotinine in the urine of the nine smokers was 16.5 ± 1.2 h. The coefficients of variation between subjects was 21 percent. Nonsmokers who were exposed to ETS had a mean $t_{1/2}\beta$ for cotinine elimination of 27.3 ± 1.9 h. The correlation coefficient for goodness of fit was r = 0.89 ± 0.02 in smokers and r = 0.90 ± 0.01 for nonsmokers exposed to ETS. For both smokers and nonsmokers, no correlation was observed between peak cotinine levels and elimination time. Nonsmokers had a substantially longer mean $t_{1/2}\beta$ of cotinine elimination (27.3) than smokers who quit (16.5).

Discussion

The results of this study support the concept that cotinine elimination can be more rapid in smokers than in nonsmokers who are exposed to ETS, and the earlier work by Kyerematen, *et al*, who observed that the $t_{1/2}\beta$ of nicotine in the urine of nonsmokers was more than twice as long as that of smokers.[17] Taken together, these observations could reflect a heightened carcinogenic potential of ETS in nonsmokers.

The slower clearance in nonsmokers of a terminal nicotine metabolite suggests that other tobacco smoke constituents or their metabolites may also remain in the bodies of nonsmokers for longer periods of time. Conceivably, this could increase the potential for endogenous formation of carcinogenic tobacco specific nitrosamines.[5] In considering

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TABLE 1—Cotinine Elimination in Urine of Smokers and ETS-Exposed Nonsmokers

Subject	$t^{1}_{2\beta}$ (h)	r	Y intercept (ng/mg creatinine)	Peak Cotinine Levels (ng/mg creatinine)
Smokers				
A	15.3	0.80	2568	7317
B	13.6	0.95	167	268
C	12.1	0.92	1388	1594
D	18.7	0.89	3006	2999
E	17.4	0.96	345	1522
K	17.5	0.86	905	839
L	17.6	0.84	48	72
M	23.3	0.81	215	207
N	13.5	0.94	1417	2358
Mean ± (SEM) 16.5 ± 1.2 (95% CI = 13.6, 18.6)				
Nonsmokers				
F	29.7	0.89	29	25
G	23.3	0.95	62	95
H	19.9	0.94	14	34
I	33.4	0.67	20	64
J	21.4	0.88	24	66
O	18.8	0.96	43	58
P	32.3	0.94	41	91
Q	31.6	0.87	33	31
R	34.0	0.90	112	117
S	28.8	0.81	19	24
Mean ± (SEM) 27.3 ± 1.9 (95% CI = 25.3, 29.4)[a,b]				

[a]Significantly different from smoker group ($p < 0.0007$).
[b]The confidence interval for the difference between means (10.8 h) was determined by first obtaining a pooled estimate of the variance.[26] Since both confidence limits were positive (95% CI = 1.4, 20.2), we concluded that the means between the groups were different.

the toxic burden imposed by exposure to cigarette smoke, it is important to note that sidestream smoke which is a major fraction of ETS contains higher levels of certain toxic cocarcinogenic and cancerogenic compounds than does mainstream smoke.[18] Before dilution with room air, some of these compounds are 50- to 100-fold higher in the sidestream smoke compared to their concentration in mainstream smoke.[19,20]

It is well documented that many compounds contained in tobacco smoke induce hepatic and pulmonary mixed function oxidases, yet the magnitude and duration of induction remains unclear.[18,21] It is conceivable that inhaled mainstream smoke and ETS might have varying effects on liver or lung oxidases due to the varying concentrations of smoke constituents. The importance of lung metabolism of inhaled xenobiotics is recognized.[22]

Differences between smokers and nonsmokers in the metabolism of drugs are substantial. Early studies by Hart, *et al*, have demonstrated enhanced drug metabolism in smokers by using anti-pyrine as an index of hepatic metabolism.[23] Powell, *et al*, reported a theophylline half-life intermediate between smokers and nonsmokers in exsmokers who abstained for three years.[24] Thus, differential rates of metabolism of nicotine in smokers and nonsmokers might be predicted from earlier work.

In our study, the $t_{1/2\beta}$ in never-smokers exposed to ETS under strictly controlled conditions was significantly longer than that seen for smokers. In most cases, the peak cotinine levels for smokers were much higher than in the ETS-exposed subjects, but no relation existed between peak cotinine levels and the $t_{1/2\beta}$ for cotinine excretion. It is probable, therefore, that elimination rate is not as greatly affected by dose as by method of exposure or subject metabolism.

A recent study by Jarvis, *et al*, concluded that differences in cotinine elimination do not exist between smokers and neversmokers.[25] In that study, nicotine pellets were ingested by five subjects of mixed smoking habits (cigars, cigarettes, and ex-smokers). Since such a protocol fails to allow for possible metabolic differences between smokers and never-smokers, ignores the lung as a metabolic site, and precludes the possibility of synergestic effects of any of the more than 3,800 constituents in tobacco smoke, we feel that the data of Jarvis cannot be compared to the current controlled study of cotinine elimination following active smoking or ETS exposure in smokers and never-smokers.

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Use of Smokeless Tobacco, Cigarette Smoking, and Hypercholesterolemia

LARRY A. TUCKER, PHD

Abstract: The primary purpose of this study was to determine the extent to which regular use of smokeless tobacco is associated with hypercholesterolemia (≥ 6.2 mmol/L) among 2,840 adult males. The confounding effects of age, education, physical fitness, body fatness, and other tobacco use were also examined. After adjustment, smokeless tobacco users were 2.5 times, heavy smokers were 2 times and mild/moderate smokers were 1.5 times more likely to have hypercholesterolemia than non-users of tobacco. Cigarette smokers did not differ significantly from users of smokeless tobacco regarding hypercholesterolemia. Users of smokeless tobacco were younger and less educated compared to non-users of tobacco, while smokers were less educated and less physically fit. (*Am J Public Health* 1989; 79:1048–1050.)

Introduction

National estimates indicate that at least 12 million Americans used some form of smokeless tobacco during 1985 and recent data show that 16 percent of males 12 to 25 years of age have used smokeless tobacco within the past year. An estimated 6 million persons use smokeless tobacco at least weekly and rates seem to be increasing, especially among adolescent and young adult males.[1]

The increased appeal and use of smokeless tobacco has generated considerable public health concern because research indicates that snuff and chewing tobacco can be significant health hazards. Most notably, dipping and chewing have been linked to oral cancer,[1–5] and numerous clinical studies have shown strong associations between smokeless tobacco use and noncancerous and precancerous oral conditions.[6–8]

Many of the health problems associated with tobacco use are a consequence of nicotine. Since the blood nicotine levels which result from smokeless tobacco use are similar to those from cigarette smoking,[9–11] the nicotine-related health effects of smoking would also be expected to result from using smokeless tobacco. Compared to nonsmokers, smokers tend to have elevated levels of low density and very low density lipoproteins and reduced levels of high density lipoproteins,[12–17] a lipid profile strongly associated with increased risk of atherosclerosis and coronary heart disease.[18–22] To date, only one unpublished study has been conducted to determine the effects of smokeless tobacco use on lipid metabolism.[23]

The present study was conducted to determine the extent to which use of smokeless tobacco contributes to hypercholesterolemia controlling for lifestyle and demographic factors, and to compare the effects of smokeless tobacco and cigarettes on hypercholesterolemia.

Methods

Subjects

A sample of 2,840 adult males with a mean age of 40.7 (SD = 10.8) was studied. Subjects were employees of over 25 different companies that participated in the Health Examination Program offered by Health Advancement Services (HAS), Inc. Approximately 70 percent of the men were married, 78 percent were White, and 73 percent had some college education. The median and modal annual gross family income was $25,000–$30,000.

All data were collected by registered nurses employed by HAS, Inc. Each subject was examined individually and privately for approximately 60 minutes after participating in an orientation and completing an informed consent form.

Instrumentation and Procedures

A written questionnaire was administered to assess demographic and life-style information, including use of smokeless tobacco and cigarettes. A Harpenden skinfold caliper was employed to assess subcutaneous fat at three body sites and the sum of the skinfold measurements along with age and gender were used to calculate the total body fat percentage of each subject.[24] Physical fitness was assessed using a step test, the Kasch 3-minute Pulse Recovery Test.[25] Approximately 10cc of blood was drawn from each subject and analyzed using the enzymatic method to determine serum cholesterol levels.[26]

Data Analysis

Subjects were classified as regular users of smokeless tobacco, mild/moderate smokers (1–20 cigarettes/day), heavy smokers (>20 cigarettes/day) or non-users of tobacco, according to their questionnaire responses. High blood cholesterol or hypercholesterolemia was defined as a total serum cholesterol level of 6.2 mmol/L or greater, consistent with the criteria of the National Cholesterol Education Program New Cholesterol Adult Treatment Guidelines.[27] The control variables were categorized as depicted in Table 1.

The associations between smokeless tobacco use, cigarette smoking, and hypercholesterolemia were measured by the odds ratio.[28] To control for potential confounders, Man-

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TABLE 1—Smokeless Tobacco Use, Cigarette Smoking, and Hypercholesterolemia According to the Control Variables

Control Variables	col %	Tobacco Use				Hypercholesterolemia
		None	Smoke-less	1–20 cig	> 20 cig	6.2 + mmol/L
Total Group	100.0	76.7	3.3	15.1	4.9	18.6
Age (years)						
19–29	15.8	74.6 (6.1)	6.6 (13.8)	16.1 (2.8)	2.7 (16.7)	6.3
30–39	34.8	76.9 (13.5)	3.3 (9.4)	14.4 (21.3)	5.5 (14.8)	14.6
40–49	27.1	77.0 (20.7)	2.5 (31.6)	14.5 (33.6)	6.0 (37.0)	23.7
50+	22.3	77.1 (27.0)	1.9 (33.3)	17.1 (29.0)	3.8 (29.2)	27.9
Body Fat						
lean	13.5	75.3 (3.8)	4.5 (0.0)	14.7 (8.7)	5.5 (11.8)	4.8
moderate	58.4	77.7 (15.7)	3.2 (18.2)	14.6 (24.4)	4.5 (21.3)	17.4
obese	28.1	75.7 (29.2)	3.2 (33.3)	15.1 (34.7)	6.0 (33.3)	30.4
Education						
elem/hs	22.2	60.6 (17.3)	6.5 (18.4)	24.6 (21.0)	8.3 (27.1)	19.3
trade/voc	4.6	62.7 (13.5)	5.9 (14.3)	20.3 (37.5)	11.0 (23.1)	19.0
college	38.1	79.4 (15.1)	3.0 (16.7)	11.7 (25.0)	5.9 (15.5)	16.4
grad/prof	35.1	87.3 (18.9)	.9 (25.0)	9.6 (19.3)	2.2 (45.0)	19.8
Fitness						
good	21.7	87.1 (11.1)	2.9 (17.7)	8.0 (19.4)	2.0 (7.7)	12.0
average	30.2	80.4 (17.5)	3.6 (17.2)	11.4 (24.4)	4.7 (23.1)	18.9
poor	19.3	76.1 (23.9)	4.0 (9.5)	14.2 (28.6)	5.7 (25.0)	24.3
did not take	28.8	75.0 (17.6)	3.2 (29.2)	15.2 (20.7)	6.6 (28.9)	19.5

NOTE: Values in parentheses reflect the percentage of subjects within the subgroup who had hypercholesterolemia. For example, 33.3% of the subjects who were 50 years of age or older who used smokeless tobacco had hypercholesterolemia.
Body fat: lean = 10% or less body fat; moderate = 11–19% body fat; obese = 20% + body fat
Fitness: (age < 27) good = recovery HR ≤ 83 bpm; average = rec HR 84–99 bpm; poor = rec HR 100 + bpm
(age ≥ 27) good = recovery HR ≤ 87 bpm; average = rec HR 88-107 bpm; poor = rec HR 108 + bpm
didn't take = those subjects who chose not to take the step test.

tel-Haenszel summary risk estimates were employed.[29–31] Analysis of covariance (ANCOVA) for unbalanced data was employed to determine mean cholesterol differences between the groups after adjusting for age, education, fitness, and additional tobacco use.

Results

Of the 2,840 subjects, 93 (3.3 percent) were regular users of smokeless tobacco, 568 (20 percent) were cigarette smokers, and 10 (.35 percent) subjects reported regular use of both products. The serum cholesterol global mean was 5.30 mmol/L (SD = 1.08). As revealed in Table 1, regular use of smokeless tobacco was reported more often by the younger and the less educated subjects, and cigarette smoking was more prevalent among the less educated and the less fit. Hypercholesterolemia, measured in nearly 20 percent of the subjects, was more common in older, fatter, and less fit subjects.

Table 2 shows the estimated relative risk of hypercholesterolemia by tobacco use, including smokeless tobacco, without adjustment and with adjustment for age, education, physical fitness, and additional tobacco use. Control for differences in body fat had little effect on the results.

TABLE 2—Estimated Risk of Hypercholesterolemia by Use of Tobacco

Use of Tobacco	Variable Controlled	Hypercholesterolemia			
		N	%	RR_{mh}	95% CI
no use (n = 2,179)	none	377	17.3	1.00	—
smokeless tobacco (n = 93)	none	17	18.3	1.08	0.63, 1.85
	age, educ, fitness, & smoking			2.51	1.47, 4.29
cigarettes 1–20 daily (n = 429)	none	100	23.3	1.45	1.13, 1.86
	age, educ, fitness, & smokeless use			1.51	1.14, 2.00
cigarettes > 20 daily (n = 139)	none	34	24.5	1.55	1.04, 2.31
	age, educ, fitness, & smokeless use			1.98	1.29, 3.03

NOTE: Subjects who reported no use of tobacco were used as the reference group.

After adjustment, subjects who used smokeless tobacco regularly displayed 2 1/2 times the prevalence of hypercholesterolemia compared to non-users of tobacco. Heavy smokers had twice the prevalence, and light/moderate smokers had 1 1/2 times the prevalence of elevated cholesterol after controlling for the potential confounders (for the smokers, use of smokeless tobacco was controlled rather than cigarette smoking). There were no differences in risk of hypercholesterolemia between the smokeless tobacco and cigarette smoking groups.

Adjusted serum cholesterol means for the non-users of tobacco, smokeless tobacco users, light/moderate smokers, and heavy smokers were 5.29, 5.36, 5.38 and 5.52 mmol/L, respectively.

Discussion

Why smokeless tobacco users had 2 1/2 times the prevalence of hypercholesterolemia compared to non-users of tobacco is not clear. The substantial proportion of hypercholesterolemia among smokeless tobacco users could be the function of multiple factors and may not reflect cause and effect. Selection bias is possible, although unlikely (i.e., those with high blood cholesterol may choose to use smokeless tobacco more than the norm).

Other potential confounders, notably diet, could account for the smokeless tobacco/hypercholesterolemia relation. Moreover, because subjects in this study had relatively high socioeconomic status and educational levels, generalization to poorer, less educated groups will require additional study. The association between smokeless tobacco use and hypercholesterolemia is biologically sound, however, since smokeless tobacco acts as a vehicle to deliver nicotine to the system and results in blood nicotine levels similar to those attained by cigarette smoking.[1,11] Hence, it is plausible that increased levels of blood cholesterol result from dipping and chewing.

In conclusion, it is apparent from the present findings that the consequences of using smokeless tobacco may reach beyond the oral cavity. Although cause-and-effect conclusions are premature given the design of this study, carrying "a pinch between the cheek and gum" may lead to hypercholesterolemia and ultimately cardiovascular disease, as well as leukoplakias and oral cancer. Clearly, increased efforts directed toward elucidating the non-oral health effects of using smokeless tobacco are warranted.

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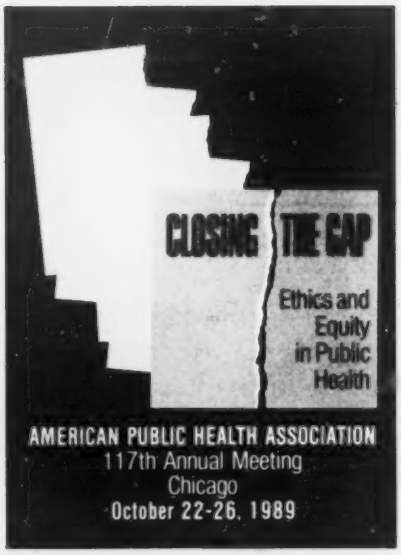

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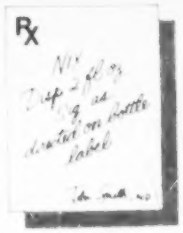

PEDICULICIDAL/OVICIDAL ACTIVITIES: *In vitro* data indicate that permethrin has pediculicidal and ovicidal activity against *Pediculus humanus* var. *capitis*. The high cure rate (97-99%) of Nix in patients with head lice demonstrated at 14 days following a single application is attributable to a combination of its pediculicidal and ovicidal activities and its residual persistence on the hair which may also prevent reinfestation.

INDICATIONS AND USAGE: Nix is indicated for the single-application treatment of infestation with *Pediculus humanus* var. *capitis* (the head louse) and its nits (eggs). Retreatment for recurrences is required in less than 1% of patients since the ovicidal activity may be supplemented by residual persistence in the hair. If live lice are observed after at least seven days following the initial application, a second application can be given.

CONTRAINDICATIONS: Nix is contraindicated in patients with known hypersensitivity to any of its components, to any synthetic pyrethroid or pyrethrin, or to chrysanthemums.

WARNING: If hypersensitivity to Nix occurs, discontinue use.

PRECAUTIONS:

General: Head lice infestation is often accompanied by pruritus, erythema, and edema. Treatment with Nix may temporarily exacerbate these conditions.

Information for Patients: Patients with head lice should be advised that itching, redness, or swelling of the scalp may occur after application of Nix. If irritation persists, they should consult their physician. Nix is not irritating to the eyes; however, patients should be advised to avoid contact with eyes during application and to flush with water immediately if Nix gets in the eyes. In order to prevent accidental ingestion by children, the remaining contents of Nix should be discarded after use.

Combing of nits following treatment with Nix is not necessary for effective treatment. However, patients may do so for cosmetic or other reasons. The nits are easily combed from the hair treated with Nix after drying.

Carcinogenesis, Mutagenesis, Impairment of Fertility: Six carcinogenicity bioassays were evaluated with permethrin, three each in rats and mice. No tumorigenicity was seen in the rat studies. However, species-specific increases in pulmonary adenomas, a common benign tumor of mice of high spontaneous background incidence, were seen in the three mouse studies. In one of these studies there was an increased incidence of pulmonary alveolar-cell carcinomas and benign liver adenomas only in female mice when permethrin was given in their food at a concentration of 5000 ppm. Mutagenicity assays, which give useful correlative data for interpreting results from carcinogenicity bioassays in rodents, were negative. Permethrin showed no evidence of mutagenic potential in a battery of *in vitro* and *in vivo* genetic toxicity studies. Permethrin did not have any adverse effect on reproductive function at a dose of 180 mg/kg/day orally in a three-generation rat study.

Pregnancy: *Teratogenic Effects*: Pregnancy Category B. Reproduction studies have been performed in mice, rats, and rabbits (200-400 mg/kg/day orally) and have revealed no evidence of impaired fertility or harm to the fetus due to permethrin. There are, however, no adequate and well-controlled studies in pregnant women. Because animal reproduction studies are not always predictive of human response, this drug should be used during pregnancy only if clearly needed.

Nursing Mothers: It is not known whether this drug is excreted in human milk. Because many drugs are excreted in human milk and because of the evidence for tumorigenic potential of permethrin in animal studies, consideration should be given to discontinuing nursing temporarily or withholding the drug while the mother is nursing.

Pediatric Use: Nix is safe and effective in children two years of age and older. Safety and effectiveness in children less than two years of age have not been established.

ADVERSE REACTIONS: The most frequent adverse reaction to Nix is pruritus. This is usually a consequence of head lice infestation itself, but may be temporarily aggravated following treatment with Nix. 5.9% of patients in clinical studies experienced mild temporary itching; 3.4% experienced mild transient burning/stinging, tingling, numbness, or scalp discomfort; and 2.1% experienced mild transient erythema, edema, or rash of the scalp.

DOSAGE AND ADMINISTRATION:

Adults and Children: Nix is intended for use after the hair has been washed with shampoo, rinsed with water and towel dried. Apply a sufficient volume of Nix to saturate the hair and scalp (especially behind the ears and on nape of the neck). Nix should remain on the hair for 10 minutes before being rinsed off with water. A single treatment is sufficient to eliminate head lice infestation. Combing of nits is not required for therapeutic efficacy, but may be done for cosmetic reasons or to meet school "no nit" policies. A nit comb is provided.

SHAKE WELL BEFORE USING.

HOW SUPPLIED: Nix (Permethrin) 1% (wt./wt.) Creme Rinse is supplied in plastic squeeze bottles that contain 2 fl. oz., weighing 56 g. (NDC-0081-0780-81).

Store at 15°-25°C (59°-77°F).

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The Impact of Medicaid on Physician Use by Low-Income Children

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Abstract: This study evaluated the determinants of physician use by low-income children, with an emphasis on the effect of Medicaid. Data are from the 1980 National Medical Care Utilization and Expenditure Survey. Regression analysis revealed that Medicaid children were more likely than both privately insured and uninsured children to visit an office-based physician. Also, Medicaid children with at least one visit to any setting had a higher number of visits than uninsured children. Such factors as age, health status, number of children in the family, educational status, and income also accounted for differences within the low-income population. The results suggest that access to physicians' services (including office-based physicians) can be increased by expanding Medicaid eligibility to uninsured low-income children and by improving private health insurance benefits among the underinsured. (*Am J Public Health* 1989; 79:1220–1226.)

Introduction

Children under age 21 comprise the single largest beneficiary group covered by the Medicaid program.[1] Congress has recognized the importance of this program by expanding Medicaid eligibility to previously uncovered groups of pregnant women and children under the Deficit Reduction Act of 1984, the Omnibus Budget Reconciliation Acts of 1986 and 1987, and most recently the Medicare Catastrophic Act of 1988. Whereas in 1964, 33 percent of the poor children and 15 percent of the nonpoor children had no physician contact in the previous two years, by 1981 the figures had decreased to 11 percent (poor) and 10 percent (nonpoor).[2,*] Average use by poor children increased from 2.3 contacts in 1964 to 4.8 contacts in 1981; nonpoor children went from 4.0 to 4.1 contacts.[2,*]

Not all of this large increase in physician use by poor children can be attributed to the Medicaid program, because of the concurrent efforts to stimulate the supply of services in low-income areas.[3] Moreover, it is misleading and inaccurate to assume that the low-income population is synonymous with the Medicaid population. For example, in 1980, 46 percent of the low-income children (less than 150 percent of poverty level) had Medicaid coverage for at least part of the year; 38 percent had private insurance for all or part of the year; and 16 percent were uninsured for the entire year.[4]

Previous research related to the effect of Medicaid on pediatric utilization is suggestive that Medicaid improves access to health care among low-income children, while simultaneously limiting their access to office-based physicians. Gortmaker concluded that Medicaid perpetuated a "two-tiered system" of care for children in Flint, Michigan.[5] Colle and Grossman support Gortmaker's view.[6]

Ongoing child health surveys in Rochester, New York, provide additional evidence on the effect of Medicaid on a child's regular source of care.[7–10] After the implementation of Medicaid, a substantial shift in source of care occurred in Rochester: from a physician's office to a clinic or health center.[9] Controlling for a wide range of factors associated with utilization, Wolfe[10] found that children on Medicaid were less likely to visit private physicians and more likely to use health centers and clinics in that city.

Descriptive data[2] clearly indicate that low-income children have experienced an increase in physician contacts since the implementation of Medicaid. However, given the sizable proportion of poor children not on Medicaid, aggregate data on the low-income population may mask utilization differences between those on Medicaid and those not on Medicaid.

This study was designed to evaluate the determinants of physician utilization by low-income children in the United States in 1980. Special emphasis has been given to the effect of Medicaid coverage on the probability and volume of use, as well as on the place of visit—physician visits in any setting versus those in a physician's office. A previous monograph[4] published by the National Center for Health Statistics presented descriptive data on the characteristics of low-income children and their health service use. This article presents the multivariate analysis, which determines the *independent* effect of many factors on the probability and volume of use.

Methods

This study uses data from the National Household Survey component of the National Medical Care Utilization and Expenditure Survey (NMCUES), a panel study of the United States noninstitutionalized civilian population in 1980. NMCUES gathered data on health status, accessibility and utilization of health services, charges and sources of payment, and health insurance coverage. The household survey consisted of 17,123 persons from about 6,000 randomly selected households. Each household was interviewed five times during 1980 and early 1981, roughly every three months. Methods have been described in detail elsewhere.[4]

The sample for this study includes children under age 18 living in families with incomes below 150 percent of the federal poverty level. Only children who were eligible for the survey for the entire survey period were included. This criterion excludes children who were born or who died during 1980, as well as those who were institutionalized for part of the year. Of the 5,074 children ages one through 17 who were eligible for NMCUES for the entire survey period, 1,409 were living in families below 150 percent of the federal poverty level. Of these, 654 had Medicaid coverage for all or part of the year.

Secondary data were obtained from two sources: 1) the Area Resource File[11] provided county level data on the supply of primary care physicians and the availability of emergency rooms; 2) a state-level exogenous price measure is based on work by the Urban Institute, to reflect the relative reimbursement within states under Medicaid and Medicare.[12]

The complex stratified multistage probability sample design used in the NMCUES required adjustments both in terms of sample weighting and variance estimation. All results have been weighted to produce national estimates.

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*Data for 1981 were obtained from updated supplemental tables to Reference 2.

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Variances have been estimated using the Taylor series approximation. The SESUDAAN computer software[13] produced standard errors for means, totals, and proportions with adjustments for the complex sample design. The SURREGR program[14] handled regression estimation for continuous dependent variables. Standard errors for logistic regression were adjusted using a method developed by researchers at the National Center for Health Services Research.**

Dependent Variables

This study examines the effect of Medicaid on four measures of physician use: the probability and the volume of physician visits in an office setting and in any setting. The effect of Medicaid on physician access is expected to vary according to locus of care.[6,10]

The distribution of physician use is non-normal due to a concentration of those with no visits at one extreme and a small number of high users at the other extreme. The approach used in this analysis is to treat the use/nonuse decision as a dichotomous choice, based on the cumulative logistic probability function which predicts the probability of seeking care within the (0,1) range.[15] The volume equations include only children with one or more visits and are handled by weighted least squares regression, since the dependent variables are continuous measures. Because of the non-normal distribution of visits among users, the volume measures are log-transformed. Thus, the probability equation reflects the decision to initiate contact, while the volume equation captures the decision to make a single or multiple visits, a decision which involves the provider as well as the patient.[16]

For the purpose of this analysis a physician visit is defined as a face-to-face contact with a physician or a nonphysician working under the supervision of a physician. In addition, visits to nurse practitioners and physician assistants who were reported as "independent providers" are included. Otherwise, visits to independent providers (primarily chiropractors and optometrists) are excluded. Mental health visits are also excluded, as defined by the condition (mental disorder), the provider (psychiatrist, psychologist, social worker), or the setting (psychiatric clinic). Telephone contacts and physician visits to hospital inpatients are not counted.

Independent Variables

The first group of independent variables includes sociodemographic characteristics of the child and the family. Use decreases from infancy through the school-age years and then increases during adolescence.[5,10,17–20] To capture the nonlinear effects, age is specified in linear and squared form. Race is specified as a dummy variable equal to one if the child is non-White. Number of children in the family and educational status of the family (a dummy variable equal to one if no member of the family graduated high school) are also included in the model.

Income is specified as actual (reported) income in 1980. Price is defined as the out-of-pocket charge relative to the total charge for physician visits; however, this measure is restricted to the volume of use equations. An alternate proxy for "price" is the type of insurance, a measure of the effect of third-party coverage. Two dummy variables are used: private insurance for all or part of the year; and no insurance coverage during the entire year. The omitted category is Medicaid coverage during all or part of the year. Thus, the coefficients on the insurance variables reflect the difference between the specified type of coverage and Medicaid coverage (as measured by the intercept).

A regular source of care has been considered the second most powerful determinant of use, following measures of health status.[21] Because those who are high users of medical care may have a regular source of care precisely due to their high use, the probability equations do not control for presence of a regular source. The volume equations, however, test the effect of a regular source on the number of visits among those with at least one. A distinction is made between a physician's office as a regular source and all other settings. (The omitted category is "no regular source.")

Three variables are included in this study to capture provider availability:

- aggregate physician stock is measured as the ratio of office-based patient care physicians engaged in primary care, per 100,000 persons in the child's county of residence;
- the number of emergency rooms relative to the county's land area is considered a proxy for emergency room access; and
- a fee ratio measures the relative generosity of reimbursement levels under Medicaid and Medicare for a general practitioner.[12]

Previous research has shown that Medicaid participation is sensitive to average charges in the private market, as proxied by Medicare reimbursement.[22,23] Average Medicare reimbursement would no longer be an adequate proxy for changes in the private market because of a recent freeze on physician charge updates and continued erosion of Medicare fee inflation due to limits imposed by the Medicare Economic Index. However, the measure used here is from 1980 and may be considered a reasonable proxy for this study.

Three measures of health status are specified in the current study: 1) a combined measure of perceived health status and activity limitation at the beginning of the year; 2) the number of bed-days during the year; and 3) a hospitalization during 1980. This last measure is expected to be associated with more ambulatory care.[6,10,17]

Results

Descriptive Results

About one-fourth of the 63.9 million US children under age 18 in 1980 lived in poor and near poor families, defined as

TABLE 1—Insurance Coverage of Low-Income Children by Poverty Status: United States, 1980

Insurance Coverage	Number*	Percent
Total	16,846	100.0
Medicaid	7,726	45.9
Full year	5,264	31.2
Part year	515	3.1
With private insurance	1,947	11.6
Private insurance	6,425	38.1
Full year	5,011	29.7
Part year	1,414	8.4
No insurance	2,695	16.0

*Numbers in thousands.
NOTE: Low-income children live in families with incomes below 150 percent of the federal poverty level.
SOURCE: 1980 National Medical Care Utilization and Expenditure Survey.

**Cohen SB, Gridley G: Present limitations in the availability of statistical packages for the analysis of complex survey data. Paper presented at the annual meeting of the American Statistical Association, 1981.

TABLE 2—Physician Visits, According to Type of Visit and Insurance Coverage: Low-Income Children, United States, 1980

Insurance Coverage	No Physician Visit in 1980		Physician Visits per Child		Physician Visits per Child with Visit	
	Any Setting	Office Setting	Any Setting	Office Setting	Any Setting	Office Setting
	Percent of Total		Average Number			
Total	29.0	52.2	2.7	1.3	3.8	2.7
Medicaid	24.8	49.2	2.9	1.4	3.9	2.8
Private insurance	31.0	55.5	2.8	1.2	4.1	2.7
No insurance	36.3	52.9	1.8	1.1	2.8	2.3
			Standard Errors			
Total			0.16	0.07	0.18	0.10
Medicaid			0.15	0.10	0.17	0.16
Private insurance			0.33	0.12	0.40	0.19
No insurance			0.19	0.14	0.20	0.18

SOURCE: 1980 National Medical Care Utilization and Expenditure Survey.

families with incomes below 150 percent of the federal poverty level. Nearly one-half (46 percent) of the 16.8 million low-income children were covered by Medicaid for all or part of 1980; another 38 percent were privately insured for all or part of the year; and 16 percent were uninsured the entire year (Table 1). When added to the 3 percent with part year Medicaid coverage and 8 percent with private insurance part of the year, over one-fourth were uninsured for at least part of 1980.

Uninsured children were less likely than Medicaid and privately insured children to have a physician visit (Table 2), and averaged one visit less during the year than the other two groups. In addition, uninsured children with at least one visit had significantly fewer visits than each of the other two comparable groups. About one-half of low-income children had no visits to an office-based physician in 1980; among those with any visits, the average was 2.7 visits. No striking differences are noted by type of insurance coverage.

Extrapolating from reported charges, it is estimated that $1.4 billion was spent for physician visits by low-income children in 1980. The average charge per child with a charge was $150 for those on Medicaid part of the year or privately insured, but only $64 for the uninsured. The children on Medicaid the full year averaged $118. The average charge per visit was in the range of $34 to $38, except for the uninsured who averaged $24 per visit. Clearly, subsidized care accounts for a large amount of this difference. As one would expect, the reported average out-of-pocket expense for non-Medicaid children ($42) was more than triple the average expense for Medicaid children ($13). However, those on Medicaid part of the year had significantly higher out-of-pocket expenses than those covered by Medicaid the full year ($27 versus $10). This may be due in part to the "spend-down" requirements to qualify for a state's medically needy program under Medicaid.

Multivariate Results

The Appendix defines the variables included in the regression analysis and presents means and standard errors. Table 3 presents the econometric results for the probability equations and Table 4 shows the volume equations.

Insurance coverage has an interesting effect on the probability and volume of use. The coefficients on the two dummy variables for private or no insurance coverage reflect the probability or volume of use *relative* to that under Medicaid (the omitted category). Two results are noteworthy. First, Medicaid children had a higher *probability* of visiting an office-based physician than privately insured children and uninsured children (Table 3, equation 2). However, there were no differences according to insurance coverage in the *number*

TABLE 3—Probability of Visit: Maximum Likelihood Dichotomous Logit Equations (t-values in parentheses)

Variables	Any Setting (Equation 1)	Office Setting (Equation 2)
Age	−0.235	−0.132
	(3.94)	(2.69)
Age-squared	0.010	0.007
	(2.88)	(2.37)
Race	−0.173	−0.329
	(0.82)	(1.98)
Educational status	−0.299	−0.203
	(1.43)	(1.24)
Number of children	−0.177	−0.233
	(3.03)	(4.80)
Income	−6.7E-5	−1.5E-5
	(2.67)	(0.72)
Private insurance coverage	−0.368	−0.610
	(1.81)	(3.01)
Uninsured	−0.562	−0.502
	(1.90)	(2.01)
Physicians per capita	−3.0E-4	−6.0E-4
	(1.14)	(2.84)
Medicaid generosity	0.105	0.565
	(0.28)	(1.40)
Emergency rooms per square mile	1.450	−6.4E-5
	(3.62)	(0.17)
Fair/poor health status or activity limitation	0.629	0.498
	(2.10)	(2.38)
Bed-days (log)	0.595	0.459
	(5.41)	(5.79)
Hospitalization in 1980	1.205	0.431
	(1.90)	(1.46)
Constant	2.699	0.733
	(6.14)	(1.83)
Adjusted R-Squared	0.267	0.090
n	1,363	1,363
Chi-Squared	464.30	197.56

NOTES: 1) The t-values were calculated after adjustments for the complex sample design. The critical t-ratio at the 5 percent level of significance is 1.96 for a two-tailed test.
2) E(integer) signifies that the value should be multiplied by the indicated power of ten.
3) Regression coefficients presented here are unstandardized parameters.
SOURCE: 1980 National Medical Care Utilization and Expenditure Survey.

TABLE 4—Number of Visits among Those with at Least One: Weighted Least Squares Regression Equations (t-values in parentheses)

Variables	Any Setting		Office Setting	
	Equation 3	Equation 4	Equation 5	Equation 6
Age	−0.115	−0.112	−0.102	−0.101
	(7.51)	(7.20)	(5.78)	(5.58)
Age-squared	0.006	0.006	0.006	0.006
	(6.40)	(5.99)	(5.80)	(5.62)
Race	−0.069	−0.075	−0.062	−0.072
	(1.10)	(1.25)	(0.90)	(1.08)
Educational status	−0.169	−0.150	−0.116	−0.119
	(3.37)	(3.10)	(2.08)	(2.09)
Number of children	−0.051	−0.046	−0.043	−0.037
	(3.62)	(2.93)	(2.49)	(2.19)
Income	−1.8E-5	−2.1E-5	2.8E-6	1.8E-6
	(2.64)	(2.75)	(0.32)	(0.20)
Out-of-pocket payments	−0.017	—	−0.012	—
	(4.47)		(3.30)	
Private health insurance	—	−0.016	—	−0.094
		(0.25)		(1.54)
Uninsured	—	−0.168	—	−0.122
		(2.67)		(1.73)
Regular source: MD office	0.012	0.120	0.078	0.057
	(1.55)	(1.24)	(0.75)	(0.62)
Regular source: other setting	0.153	0.160	−0.215	−0.204
	(1.52)	(1.54)	(2.01)	(1.88)
Physicians per capita	−1.4E-4	−1.4E-5	−8.6E-5	−8.4E-5
	(1.71)	(1.74)	(1.25)	(1.29)
Medicaid generosity	−0.045	−0.066	−0.052	−0.060
	(0.32)	(0.46)	(0.37)	(0.44)
Emergency rooms per square mile	2.8E-4	2.9E-4	−9.5E-5	−8.5E-5
	(2.79)	(2.80)	(1.16)	(1.01)
Fair/poor health status or activity limitation	0.401	0.413	0.257	0.265
	(5.47)	(5.68)	(3.26)	(3.38)
Bed-days (log)	0.218	0.216	0.160	0.160
	(9.97)	(9.74)	(5.14)	(5.10)
Hospitalization in 1980	0.400	0.393	0.365	0.364
	(4.14)	(4.02)	(2.88)	(2.36)
Constant	1.359	1.362	0.978	0.977
Adjusted R-Squared	0.302	0.296	0.224	0.217
n	972	972	639	639
F	63.53	51.20	16.58	16.06

NOTES: 1) Equations (3) and (5) include a continuous measure of out-of-pocket price. Equations (4) and (6) include two dummy variables for insurance coverage (Medicaid coverage is the omitted category).
2) The t-values were calculated after adjustments for the complex sample design. The critical t-ratio at the 5 percent level of significance is 1.96 for a two-tailed test.
3) E(integer) signifies that the value should be multiplied by the indicated power of ten.
4) Regression coefficients presented here are unstandardized parameters.
SOURCE: 1980 National Medical Care Utilization and Expenditure Survey.

of visits to office-based physicians among those with at least one visit (Table 4, equation 6). The second important result is that uninsured children had fewer visits overall (any setting), compared to other low-income children covered by Medicaid or private insurance (Table 4, equation 4).

As an alternative to the insurance variables, the price effect was measured as the percent of total charges paid out-of-pocket. As out-of-pocket charges increased, physician use decreased. Both total visits (any setting) and office-based visits were negatively related to "price." Use of office-based physicians was more sensitive to price than total use, suggesting a possible quantity-quality substitution.[24] In addition, non-Medicaid children were more price sensitive than Medicaid children.[24]

Physician use by low-income children has a U-shaped age effect: controlling for other factors, the predicted number of visits among those with at least one visit was highest for children ages one and two and lowest for those between ages nine and 11. The predicted number of visits began rising at age 11 through age 17.

White children had a higher likelihood of an office visit than children of other races. However, among those who had at least one visit, the number of visits was similar. Educational status of the family was not associated with the probability of a visit, but children in "more educated" families had a higher number of visits, both to an office setting and any setting. The number of children in the family was negatively related to physician use. Income was also negatively related to the probability of any use as well as the number of visits per user.

The presence of a regular source, either a physician's office or other place, had little or no effect on the volume of visits. As would be expected, those who reported a "regular place" as their usual source (e.g., clinic, health center, emergency room) had fewer visits to a physician's office. In general, however, these findings suggest that having a regular source of care is not associated with a higher level of use among low-income children.

The availability of office-based primary care physicians was negatively related to the probability of an office visit,

possibly because physicians may locate in areas of a county that are not low-income so that counties could have a high rate of physicians per capita but low rates in inner-city areas where the low-income are concentrated.[25] Alternatively, areas with a high ratio of office-based physicians per capita also tend to have more non-office-based facilities, such as outpatient departments, community health centers, and public health clinics. The availability of hospital emergency rooms was found to be a significant factor in the probability and volume of visits (any setting). Thus, as the availability of emergency rooms increases, use increases. This suggests that emergency rooms still may be an important source of care for low-income children despite efforts to develop alternate sources of care (e.g., community health centers).

Health status measures were the single most important group of variables that determine who uses physicians' services and how much. The direction of causality between bed-days and use is not clear because physician use may lead to bed-days or bed-days may lead to physician use. A hospitalization in 1980 was not a significant determinant of the likelihood of physician use, but as might be expected the volume of use was higher for those hospitalized during the year, indicating that ambulatory care complements inpatient use.

Table 5 displays the annual predicted utilization of Medicaid and non-Medicaid children according to health status. In all cases, the probability or volume of use was higher among Medicaid than non-Medicaid children. Medicaid coverage clearly increased the likelihood of an office visit. In addition, those with Medicaid coverage had more visits than uninsured children; compared to privately insured children, however, differences are small. Health status was a major factor in the probability and volume of use in all cases, yet the probability of a physician visit for uninsured children in fair or poor health was lower than the probability for Medicaid children in excellent or good health, all other things being equal.

Discussion

The preceding analysis has shown that Medicaid coverage increases access to office-based physicians, within the low-income population. The results obtained in this study suggest that previous studies which focus on the general population may obscure the effect of Medicaid on physician use by low-income children. Another important effect of Medicaid was on the number of visits among those with at least one visit. By lowering the relative price of a visit, Medicaid children made more visits than uninsured children, but not compared to privately insured children. These results are particularly important in light of recent Congressional initiatives allowing states to expand Medicaid eligibility to low-income children through age eight with incomes up to 100 percent of the federal poverty line, and infants under age one with incomes up to 185 percent of the federal poverty line.[26] The Medicare Catastrophic Act of 1988 *requires* states to cover infants under age one with incomes at or below 75 percent of the poverty line by July 1, 1989 and 100 percent by July 1, 1990.

The results of this analysis suggest that a policy to expand Medicaid eligibility to low-income uninsured children may be expected to have a relatively low *social cost*, that is, the cost of services resulting from a higher demand. It is estimated that a 50 percent decrease in the relative price of a visit would result in a 10 percent increase in use.[24] Applying this estimate to the average use of uninsured children with at least one visit, a 50 percent reduction in the relative price would result in an increase from 2.8 to 3.1 visits, on average. The *shifted cost*, however, could be somewhat higher if services paid for out-of-pocket, by other third-party payors, or through uncompensated care are paid by the Medicaid program as a result of expanded Medicaid eligibility. The costs could be partly offset by increased use of lower-cost office-based physicians.

This study also has implications for the financial access of privately insured low-income children. Low-income children covered by private health insurance may have lower use of office-based physicians' services, in part, because of the limited scope of benefits covered by their policies. Furthermore, income, educational status, and number of children may be important factors in health service use by the privately insured due to the indirect costs associated with taking a child to the doctor. This group had much narrower benefits for ambulatory care than for inpatient care, with coverage generally for inpatient care (86 percent) and inpatient physician visits (84 percent), but less frequently for office visits (44 percent), and rarely for dental visits (18 percent).[24]

TABLE 5—Predicted Annual Use* of Physicians' Services, According to Insurance Coverage: Low-Income Children, United States, 1980

	Insurance Coverage					
	Children in Excellent/Good Health and No Activity Limitation			Children in Fair/Poor Health or With an Activity Limitation		
	Medicaid	Privately Insured	Uninsured	Medicaid	Privately Insured	Uninsured
Probability of Physician Visit						
Any Setting	.74	.70	.64	.81	.77	.71
Office Setting	.53	.40	.42	.64	.51	.54
Number of Physician Visits						
Any Setting	2.64	2.60	2.23	3.98	3.92	3.37
Office Setting	2.05	1.86	1.81	2.67	2.43	2.36

*The predicted utilization figures were derived by multiplying coefficients from regression equations by the mean values for the total sample. Regression equations controlled for demographic characteristics, health status, economic factors, community resources, and organizational variables.

**For those with at least one physician visit.

***For those with at least one visit to an office-based physician.

SOURCE: 1980 National Medical Care Utilization and Expenditure Survey.

In this context, it is not surprising that privately insured low-income children had a lower probability of visiting an office-based physician than Medicaid children. However, among those with at least one office visit, there was no difference. Those with at least one visit are most likely the children with some coverage for office visits or else had individual arrangements with a physician. The children with no coverage for office visits may be considered worse off than Medicaid children, and perhaps, in terms of financial access to office-based physicians, they may be equivalent to uninsured children. At the very least, they are underinsured. Even those with coverage for office visits but with large deductibles may not have effective coverage if the level of the deductible results in full payment for the few visits made by the majority of the children each year. These problems may be compounded for the 8 percent of the low-income children who were privately insured part of the year and uninsured the remainder of the year.

Expanding private insurance benefits for office visits may lead to lower program costs for insurers. Privately insured children made an extremely high proportion of visits to hospital outpatient departments.[4] Office visit coverage may shift the locus of care from costly hospital-based facilities to less costly office settings. Furthermore, evidence suggests that office-based physicians provide more continuous care than hospital facilities and tend to hospitalize less often.[27] As a first step, Congress is considering legislation that would mandate employers with more than five employees to provide coverage for well-baby care through age one (among other benefits).***

These results should be qualified by several caveats. One limitation of this study is that the utilization measures do not account for the necessity or appropriateness of use. For this reason, it is not clear that "more use is better." Thus, to say that Medicaid children have more visits or a higher likelihood of an office visit should not be interpreted to mean that policies should necessarily strive to equalize physician use among Medicaid and non-Medicaid children, regardless of need. Rather, the goal should be equitable access, that is, "enough care to achieve . . . 'an adequate level of health care.' "[25]

A second caveat concerns the relevance of this study to the current environment. This study uses data collected in calendar year 1980. However, since 1980, many changes have taken place in the financing and delivery of health services. Within the private insurance market, many firms have sought to contain the cost of private health insurance by reducing benefits and/or increasing copayments. Others have turned to self-insurance to avoid compliance with state mandated benefit regulations. Additionally, preferred provider arrangements have been negotiated by private insurers as cost-containment mechanisms. The early 1980s saw extreme turbulence in the public sector, with retrenchments in the welfare rolls and hence in Medicaid eligibility. By the mid-1980s, Congress sought to extend Medicaid coverage to more low-income children and pregnant women. Other maternal and child health programs have continued to suffer from budget reductions (notably the MCH block grant). In other words, the 1980s may have brought about a widening in the disparities between low-income children covered by Medicaid and those who are uninsured or privately insured. Additional research is required based on a more recent database to test the generalizability and sensitivity of the results obtained in this study.

***Minimum Health Benefits for all Workers Act of 1987 (S 1265, HR 2508).

APPENDIX

*Weighted Means and Standard Errors for the Study Sample**

Variables	Mean	Standard Error
Age at beginning of 1980	7.969	0.203
Age-squared	90.209	3.181
Race: = 1, if non-White; 0, otherwise	0.335	0.031
Educational status of family: = 1, if no high school graduate; 0, otherwise	0.410	0.029
Number of children under age 17 in family	3.144	0.100
Family income in 1980	8700.397	272.069
Percent of total charges paid out-of-pocket**	3.761	0.317
Private health insurance	0.381	0.027
Uninsured the entire year	0.160	0.017
Regular source: MD office	0.523	0.026
Regular source: Other setting	0.326	0.026
Primary care physicians per 100,000 population in county	383.909	22.716
Generosity of Medicaid fee relative to Medicare fee within state	0.741	0.016
Emergency rooms per 100,000 square miles	72.596	11.105
Fair or poor health status or activity limitation	0.095	0.010
Number of bed disability days (log)	0.959	0.040
Hospitalization in 1980	0.078	0.008

*The means and standard errors were generated by PROC SURREGR under SAS, and take into account the complex sample design.
**Includes only children with at least one physician visit.
SOURCE: 1980 National Medical Care Utilization and Expenditure Survey.

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Risk Factors Associated with Participation in the Ontario, Canada Doctors' Strike

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Abstract: To identify factors associated with participation in the 1986 Ontario, Canada doctors' strike, we surveyed 1,028 physicians; 69 percent responded, of whom 42 percent participated in the strike. Risk factors for participation included income greater than $135,000, being a surgeon or gynecologist, having previously "opted out" of the Ontario Health Insurance Plan, being professionally dissatisfied, being politically conservative, favoring political activism by physicians, holding a positive view of the social consequences of extra-billing, and perceiving family, associates, patients and the public to favor the strike.

Eighty percent of strikers, but 32 percent of non-strikers, met criteria we established for four strike-prone groups: the "economically rational," the "ideologically committed," the "professionally disaffected," and the "socially malleable." Respondents belonging to one or more of these groups were much more likely to have participated in the strike (64 percent vs 17 percent). Strategies to deal with physician militancy should address the multiplicity of motives that appeared to have influenced doctors in Ontario. (*Am J Public Health* 1989; 79:1227–1233.)

Introduction

Health care initiatives in many countries in recent years have led to the emergence of new forms of medical practice and new kinds of professional relationships, some of which have undermined the relatively unrestricted autonomy that practitioners previously enjoyed. The professional press in the United States and elsewhere has recently reported on doctors in conflict with insurance companies, hospitals, health maintenance organizations, and governmental agencies.[1–3] At least 17 times since 1960, physicians in Western countries have translated discontent with health systems into organized work slowdowns and strikes. These activities have generally been directed against governments for the purpose of extracting financial concessions, preventing undesired structural changes in health care delivery, or preserving professional independence.[4,5] Although associated with certain medical[6,7] and moral costs,[8] they have often produced definite gains for participants.

Although physicians dissatisfied with their conditions of practice have a variety of individual and collective responses available to them,[9] strikes and work slowdowns are the most visibly disruptive. Several studies have described factors that affect physician satisfaction and dissatisfaction,[10–12] but the characteristics that differentiate physicians willing to strike from those unwilling to strike have not been examined.

Health care in Ontario, Canada is financed jointly by the provincial and federal governments and is administered through the Ontario Health Insurance Plan (OHIP). Physicians are paid on a fee-for-service basis according to a schedule negotiated annually by the Ontario Ministry of Health and the Ontario Medical Association (OMA).[13] Throughout most of the OHIP's history, physicians were permitted to "opt out" of the plan, thereby gaining the option to "extra-bill" their patients at rates of their own choosing (and presumably higher than those allowed by the provincial fee schedule). This situation ended in 1984, when Parliament directed that provinces stop allowing extra-billing or else forfeit a substantial portion of their federal funding. Along with other provincial medical societies and the Canadian Medical Association, the OMA protested, on the grounds that a ban on extra-billing would infringe upon the right of physicians to negotiate a contract directly with their patients, and ultimately would undermine the quality of care, by eliminating a safety valve for occasions when the government failed to provide adequate support to the system.

When the Ontario legislature prepared to implement the federal mandate by introducing the Health Care Accessibility Act (1986), the OMA asked its members to strike. Beginning on June 12, 1986 and ending 23 days later, the strike not only failed to avert the ban on extra-billing, but was widely viewed as a public relations disaster for physicians.[14,15]

To understand how misgivings about changes in the system were transformed into action, and to identify risk factors for participation in the strike, we surveyed a broad sample of Ontario physicians. Based on expectancy-value theory[16] and prior studies of physicians in the US and Canada,[17–19] we hypothesized that they would be influenced by economic self-interest, ideologic orientation, professional work satisfaction, and perceptions of how peers, patients and others viewed the strike.

Methods

In late April 1987, we mailed a self-administered questionnaire to a probability sample of 1,170 physicians in active practice in Ontario as of June 12, 1986. Physicians were selected in a systemic fashion from the 1986 *Ontario Medical Directory*.[20] Subjects from communities with fewer than 30 physicians were oversampled to assure adequate representation of rural practitioners. In addition to identifying factors, physicians were questioned as to duration and intensity of participation in the strike; political affiliations; attitudes as measured on a series of scales dealing with government intervention, physician activism, medical altruism, professional autonomy, and work satisfaction; prior history of "opting out" and "extra-billing"; perceptions of the consequences of extra-billing, the Health Care Accessibility Act, and of the strike itself; and degree of support for the strike among co-workers, family, staff, patients and the public.

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Subjects missing more than 50 percent of the items used to form a scale were eliminated from analyses of that scale. Although the attitudes of strikers and non-strikers on Likert-type scales were initially compared using mean scale scores, scores were ultimately dichotomized about the midpoint of the scale for ease of interpretation[21]; the main results were unchanged. Complete scaling information is given in the Appendix.

Statistical Analysis

We determined strike participation rates within demographic, professional, and attitudinal subgroups of physicians. Differences were assessed with Chi-square tests; confidence intervals for differences in participation rates were calculated using standard methods.[22]

Based on the hypothesis that economics, ideology, professional satisfaction, and perceptions of how others viewed the strike would be important determinants of strike behavior and, after noting the results of the univariate analyses, we selected nine variables for entry into three logistic regression models. Model I was the full model (see Table 2 for specification of variables). In addition to the full model, two partial models were developed in order to estimate the effects on strike participation of professional characteristics along with general attitudes not directly related to the issues at hand (Model II), and of professional characteristics alone (Model III). Adjusted odds ratios and 95 percent confidence intervals were derived from the logistic coefficients and their standard errors.[23] Reweighting the regression models to adjust for oversampling of rural physicians had no significant impact upon the results.

In the third part of the analysis, we hypothesized that striking physicians would fall disproportionately into four overlapping groups: the "economically rational," the "ideologically committed," the "professionally disaffected," and the "socially malleable." We developed criteria defining membership in the groups using variables that were both theoretically important[16,24] and statistically significant in the logistic regression analysis. Finally, we computed strike participation rates among risk-group members and calculated the sensitivity and specificity of the group classification scheme for identifying strike participation in the study population.

Results

After follow-up, 712 subjects (69 percent) returned completed surveys. Respondents and non-respondents did not differ by specialty, community size, or gender. The median age of respondents (43 years) was slightly greater than that of Ontario physicians as a whole (41 years).[25]

Eighty-percent of study participants were male; 92 percent were White. Forty-six percent of subjects were in general or family practice, 11 percent in internal medicine or pediatrics, 12 percent in surgery or obstetrics-gynecology, 8 percent in psychiatry, and 23 percent in other specialties including radiology, anesthesia, dermatology, rehabilitation, and emergency medicine. Thirty-three percent of the sample worked in solo private practice, 31 percent in group practice and the remainder in public or community hospital-based practice (11 percent), academic medicine (13 percent), residency (9 percent) and various other settings including industry and penal institutions (3 percent).

The percentage of respondents participating in the strike was 42 percent. Among strike participants, 50 percent withheld services for greater than 21 days, while the rest returned to practice-as-usual after one to 21 days. Withdrawal of services was highly selective. For example, 100 percent of "striking" surgeons reported some reduction in office visits, but only 62 percent closed their offices to patients visits altogether; 76 percent stopped performing elective surgical procedures, but only 6 percent stopped making hospital and emergency room visits.

Striking and non-striking physicians were similar in age, sex, and religion, but those who were not White, whose first language was other than English or French, who had high family incomes, were married or who had children were more likely to strike (Table 1). Those not White (95 percent Asian) and those whose first language was not English or French struck at higher rates regardless of practice setting or specialty. Being married, however, was associated with a greater likelihood of striking only among physicians with family income less than $110,000 per year. Strikers and non-strikers were also distinguished by specialty and practice setting, but not by community size. Working more than 50 hours per week was associated with greater participation in the strike only among non-surgeons.

Strike participation was greater among those who considered their political thinking "right-of-center" or "far-right" (Table 1) and among physicians who identified with the Progressive Conservative Party rather than with the Liberal Party or the New Democratic Party.* Other characteristics that were associated with increased strike participation included high valuation of professional autonomy, individualistic notions of responsibility for health care, opposition to government-sponsored social programs, and agreement that physicians could "properly" participate in leafleting, demonstrations, and various kinds of strikes (Table 1).

As expected given the nature of the dispute, behaviors and attitudes related to extra-billing were strongly associated with strike participation. Physicians who had "opted out" or "considered opting out" of OHIP were more likely to strike than those who had not. Although the vast majority of respondents (86 percent) viewed preservation of the right to extra-bill as socially desirable, those who disagreed were much less likely to strike (Table 1).

Ontario Medical Association members were more likely to have struck than non-members (Table 1), as were those who "usually or always agreed" with the OMA's official positions on medical and political issues.* Physicians who generally believed that colleagues, spouses, office staff, patients, and the public were in favor of the strike were more likely to have participated in it themselves. Finally, the 24 percent of subjects who were less than "moderately satisfied" with their professional work participated in the strike more frequently.

Logistic regression provided adjusted estimates of the effects of nine selected variables on strike participation. In the full model (Model I, Table 2), behaviors and attitudes related most directly to the strike had the strongest associations: the odds of striking for physicians who had opted out of OHIP in the past, viewed extra-billing as socially beneficial, or perceived others to have supported the strike, were three to eight times higher than for those who had not. In this model, having a high family income, practicing alone or in a non-hospital-based group, and having a positive view of physician activism also were related to striking.

In Model II, which contained practice characteristics and general attitudinal factors, but excluded variables prox-

*Data available on request to authors.

TABLE 1—Strike Participation Rates among Subgroups of Physicians

Characteristics	Strike Participation Rate	Rate Difference (Highest-Lowest)	(95% CI)
Age	%	%	%
45 years or younger (n = 393)	43		
Greater than 45 years (n = 315)	40	3	(−4, +10)
Sex			
Male (n = 568)	42		
Female (n = 137)	41	1	(−8, +10)
Religion			
Protestant (n = 370)	39		
Catholic (n = 171)	43		
Jewish (n = 87)	43		
Other or none (n = 70)	51	12	(−1, +25)
Race			
White (n = 644)	40		
Other (n = 55)	55	15	(+1, +29)
First Language			
English (n = 564)	39		
French (n = 21)	43		
Other (n = 118)	52	13	(+3, +23)
Family Income			
Less than 85,000 (n = 206)	28		
$85,000–$135,000 (n = 239)	39		
Greater than $135,000 (n = 249)	55	27	(+18, +36)
Marital Status			
Married (n = 596)	44		
Unmarried (n = 109)	30	14	(+5, +23)
Family Size			
No children (n = 224)	33		
One or more children (n = 480)	46	13	(+5, +21)
Specialty			
General or Family Practice (n = 321)	42		
Internal Medicine, Pediatrics (n = 80)	32		
Surgery, Obstetrics-Gynecology (n = 86)	58		
Psychiatry (n = 56)	41		
Other (n = 158)	37	26	(+19, +33)
Practice Setting			
Solo practice (n = 228)	50		
Group practice (n = 215)	51		
Academic or hospital practice (n = 257)	25	26	(+17, +35)
Community of Practice			
Greater than 1,000,000 persons (n = 130)	36		
50,000 to 1,000,000 persons (n = 282)	45		
Less than 50,000 persons (n = 285)	41	9	(−1, +19)
Length of Work Week			
Less than or equal to 50 hours (n = 379)	35		
Greater than 50 hours (n = 329)	49	14	(+7, +21)
Position on Political Spectrum			
Liberal-moderate (n = 418)	35		
Conservative (n = 275)	52	17	(+10, +24)
Professional Autonomy			
Highly Valued (n = 559)	47		
Less highly valued (n = 149)	20	27	(+19, +35)
Medical Altruism			
Socially oriented (n = 550)	37		
Individually oriented (n = 147)	57	20	(+11, +29)
Economic-welfare Liberalism			
Liberal (n = 118)	36		
Conservative (n = 590)	43	7	(−3, +17)
View of Physician Activism			
Negative or neutral (n = 367)	25		
Positive (n = 341)	59	34	(+27, +41)
Ontario Health Insurance Plan Status			
Never considered opting out (n = 403)	29		
Thought about opting out (n = 181)	57		
Opted out (n = 104)	71	42	(+32, +52)
View of Social Benefits of Extra-billing			
Negative or neutral (n = 99)	6		
Positive (n = 609)	47	41	(+35, +47)
OMA Membership			
Member (n = 580)	46		
Non-member (n = 123)	23	23	(+15, +31)
Impression of How Others Viewed the Strike			
Appeared to be opposed or neutral (n = 465)	28		
Appeared to be in favor (n = 233)	73	45	(+38, +52)
Physician Job Satisfaction			
Moderately satisfied or better (n = 574)	40		
Less than moderately satisfied (n = 117)	50	10	(0, +20)

TABLE 2—Adjusted Associations between Physician Characteristics and Strike Participation Using Logistic Regression

Characteristics	N	Model I* Odds Ratio for Striking	Model I* (95% CI)	Model II† Odds Ratio for Striking	Model II† (95% CI)	Model III‡ Odds Ratio for Striking	Model III‡ (95% CI)
Family Income							
Less than $85,000	202	1.0		1.0		1.0	
$85,000 to $135,000	237	1.5	(.86, 2.5)	1.4	(.87, 2.2)	1.5	(.98, 2.3)
Greater than $135,000	245	2.8	(1.7, 4.8)	3.0	(1.9, 4.7)	2.8	(1.8, 4.3)
Specialty							
General/Family Practice	315	1.0		1.0		1.0	
Int. Medicine/Pediatrics	78	2.0	(.95, 4.2)	1.7	(.88, 3.3)	1.3	(.69, 2.3)
Surgery/Gynecology	88	2.0	(.98, 4.2)	2.5	(1.3, 4.5)	3.0	(1.7, 5.3)
Psychiatry	54	0.9	(.41, 2.0)	1.2	(.57, 2.3)	1.3	(.67, 2.4)
Other	156	1.4	(.75, 2.5)	1.2	(.72, 2.1)	1.3	(.83, 2.2)
Practice Setting							
Solo	223	1.0		1.0		1.0	
Group	211	1.1	(.67, 1.8)	1.0	(.66, 1.6)	1.2	(.79, 1.8)
Academic/hospital	250	0.3	(.20, .62)	0.3	(.16, .43)	0.4	(.21, .50)
Political Position							
Liberal-moderate	395	1.0		1.0		---	
Conservative	262	1.2	(.78, 1.8)	1.9	(1.3, 2.7)	---	---
Physician Activism Views							
Negative or neutral	333	1.0		1.0		---	
Positive	324	3.7	(2.4, 5.6)	4.8	(3.3, 6.9)	---	---
Professional Satisfaction							
Moderate or better	545	1.0		1.0		---	
Less than moderate	112	1.6	(.94, 2.6)	1.7	(1.04, 2.7)	---	---
Stake in the Issues (OHIP Status)							
Never considered opting out	371	1.0		---		---	
Thought about it	168	1.8	(1.1, 2.8)	---	---	---	---
Opted out prior to strike	95	3.2	(1.7, 6.0)	---	---	---	---
View of Extra-Billing							
Negative or neutral	79	1.0		---		---	
Positive	555	8.0	(2.4, 2.7)	---	---	---	---
Impression of Others' Views							
Appeared opposed/neutral	425	1.0		---		---	
Appeared in favor	209	4.8	(3.1, 7.4)	---	---	---	---

*Full model
†Model containing professional characteristics and general attitudes only.
‡Model containing professional characteristics only.

imate to the issues at hand (Table 2), high income surgical speciality, non-institutional practice, political position, and professional dissatisfaction were significantly associated with striking. Little change in the adjusted odds ratios for practice characteristics was observed when general attitudes were dropped (Model III, Table 2), except that the point estimate of the odds ratio for surgeons rose to 3.0. The full model correctly distinguished strikers from non-strikers 78 percent of the time.

Based on the expectation that strike participants would arise disproportionately from specific segments of the physician population, we tested the proposition that strikers were more likely than non-strikers to meet defined criteria as "economically rational," "ideologically committed," "professionally disaffected," or "socially malleable" (Table 3). Physicians who generally viewed doctors' strikes as "probably or definitely improper" were excluded from each of the four groups. Eighty percent of strikers, but 32 percent of non-strikers belonged to one or more of these "high risk" groups. Of 368 physicians (52 percent) belonging to one or more of these four categories, 64 percent struck, compared to just 17 percent of the 354 subjects not-so-classified.

TABLE 3— Categories of physicians at "high risk" for participation in the Ontario doctors' strike.

High-Risk Category	Criteria
Economically rational (n = 121; strike participation rate = 77%)	1. History of opting out of OHIP or considered opting out of OHIP, and 2. Family income greater than $110,000.
Ideologically committed (n = 240; strike participation rate = 68%)	1. Right-wing or right-of-center in political thinking, and 2. Progressive Conservative Party affiliation, and 3. View of extra-billing as socially beneficial.
Professionally disaffected (n = 136, strike participation rate = 65%)	1. Dissatisfaction with aspects of ability to enlist system resources for the benefit of patients, and 2. Pessimism about the future of medicine.
Socially malleable (n = 254, strike participation rate = 72%)	1. Ontario Medical Association member and 2. Perception that either colleagues, patients, or both supported the strike.

Discussion

Medical autonomy in the United States is under fire.[1,2,26] The threat comes not only from patients' changing attitudes, grounded in consumerism and litigiousness, but also from the new-found strength of third party payers and the rise of bureaucratic forms of medical practice.[27] Under these conditions, strikes by US physicians may become increasingly common. This study suggests that, contrary to public impressions, practicing physicians in Ontario are politically

heterogeneous, constituting a collection of diverse interests rather than a professional monolith. Nonetheless, many strikers shared certain common characteristics.

The announced intent of the Ontario government to ban extra-billing in May 1986 served as a focal point for the mounting discontent of many Ontario physicians. The strike drew support from nearly all regions and specialties, but failed to enlist the participation of a majority. Strikers differed from non-strikers demographically, professionally, politically, and socially. Since the survey was conducted after the strike, some of these associations may represent an effect of the decision to strike (or its covariates) upon the attitudes or characteristics in question rather than the other way around. However, many of the variables of interest (e.g, practice setting, income) are relatively stable over time and are unlikely to be influenced by the fact of strike participation.

The increased rate of strike participation among physicians who were not White was modest and might be discounted, but the association between being a native speaker of a foreign language and striking was more robust. Analysis of written comments submitted by these physicians revealed that many of them were immigrants from eastern Europe or Asia who harbored deep resentment against government involvement in medicine. British physicians expressed similar feelings.

Conservative ideology was consistently correlated with increased strike participation, whether measured by party affiliation, self-description, or agreement with conservative values. Similarly, a study of British general practitioners[28] found that doctors who provided the Medical Guild with undated resignations from the National Health Service (to be used as a threat during negotiations over reimbursement) were more likely to endorse traditional medical role conceptions, reject self-sacrificing attitudes, and oppose government involvement in medical practice.

Variables were selected for logistic regression analysis based on *a priori* hypotheses and on the strength of their simple and stratified associations with striking. The results accorded with a general formulation of expectancy-value theory,[16] in which persons weigh the consequences of action and relevant social norms in deciding how to behave. Views on extra-billing provided the most direct measure of physicians' beliefs about the consequences of the proposed legal ban. Prior OHIP status reflected physicians' stake in the issues at hand, a factor of importance in prior studies of physician activism.[28] To flesh out the model, we tried to determine subjects' prior conceptions of the likely consequences of going on strike (for the practice of medicine and for the public health), but the proposed scale was not sufficiently reliable (Cronbach alpha < 0.5).

Several factors appeared to modify the effects of perceived consequences and subjective norms in our subjects. Respondents who disapproved of political expression by physicians were unlikely to have participated in the strike, no matter what their politics. Whether these notions derived from ethical considerations or reluctance to breach authority is unclear. Practicing surgery or gynecology was associated with striking, but the odds ratio declined by approximately 50 percent as attitudes and prior OHIP status were added to the logistic equations, suggesting that their high participation rate was largely due to their more often holding attitudes associated with striking and to their having a larger economic stake in the issues.

In order to present our findings in a format useful for policy, we restructured the logistic results to produce a classification scheme that made theoretic sense. "Economically rational" physicians stood to benefit directly from the continued ability to extra-bill and had high family incomes. For them, the expected benefits of preserving extra-billing exceeded the expected costs of striking. High-income physicians could also be expected to endure a loss of practice income (as during a strike) with less hardship.

"Ideologically committed" physicians were politically conservative, identified with the Progressive Conservative Party, and considered the social benefits of extra-billing to outweigh any conceivable costs. They viewed the motives of the Ontario government with suspicion, disputed the notion that extra-billing denied the indigent or elderly access to care, and viewed extra-billing as a "safety-valve" primarily for preservation of autonomy and not just to enhance income.

The "professionally dissatisfied" doctors in our study expressed dissatisfaction with one or more aspects of their ability to provide care under the constraints of Ontario's insurance plan and were pessimistic about their professional future. They may have joined the strike as an expression of alienation from their work and as a reaction to the perceived hostility of legislators, bureaucrats, and journalists.

The final risk category included physicians who may have been influenced to strike by their medical leadership, their professional peers, or their general social environment. These "socially malleable" doctors belonged to the Ontario Medical Association and perceived colleagues or patients as favoring the strike. Membership in the OMA was included in the criteria list for "social malleability" only because *non-members* were highly *unlikely* to strike. Since we did not measure others' actual views, we cannot exclude the possibility that physicians' perceptions were biased; nor can we be sure that subjects did not select the company they kept based partly on medical politics. Nevertheless, they may have participated in the strike partly out of deference or loyalty to respected persons or groups. Given the demonstrated importance of social influence in diverse populations, it would be surprising if physicians emerged as an exception.

Although non-respondents to our survey were similar to respondents along lines of gender, specialty, and community size, we lack information about their income levels, workloads, practice settings, and political affiliations and attitudes. However, the distributions of many of these variables among respondents are remarkably similar to those reported previously in population surveys of Ontario physicians.[19,25] Furthermore, our estimate of the proportion of physicians participating in the strike jibes well with those of others.[29] Lastly, the magnitude of any differential response would have had to be extremely large to influence the direction of most of the findings. For example, even if all 316 non-respondents were in solo or group practice and only 20 percent of them struck (vs 50 percent among responding solo/group practitioners), the overall strike rate for solo/group practitioners would remain 13 percent higher than for institutionally based physicians. If, on the other hand, all 316 non-respondents practiced in institutional settings and 50 percent of them struck, the overall rate difference would still favor solo/group physicians by 12 percent.

Physicians' strikes are not inevitably harmful to society. They "can serve as an important catalyst in converting a rigid and conservative health system into a more flexible democratic organization for all its workers,"[5] and may be ethically justifiable in some situations.[30,31] Nevertheless, as quasi-violent means to even socially desirable ends, strikes frighten

the public,[7] arouse ethical qualms in physicians, have the potential for producing adverse health outcomes, and ought to be avoided if possible.

Once the federal government mandated an end to extra-billing, the Ontario strike may have been inevitable. It is possible, however, that the strike might have been attenuated by early, targeted attention to the groups of physicians described here. Some combination of fee increases, assurances of continued clinical autonomy, concrete improvements in the health care system, and more accommodating dialogue with the OMA might have reached some members of the "economically rational," "ideologically committed," "professionally disaffected," and "socially malleable" groups, respectively. In any case, this study indicates that strikers could be identified by their membership in specific risk groups, each characterized by specific attributes and motives. As physicians in the United States as well as Canada begin to exercise their collective power in defense of socio-economic and technical autonomy,[27] greater understanding of the diverse interests within the profession may facilitate more skillful management of doctors' strikes and other manifestations of this emerging medical militancy.

APPENDIX

Description of Scales

1. *Economic—Welfare Liberalism*[17]
   Internal consistency in current population: alpha = .63
   Scale: 5 = strongly agree—1 = strongly disagree
   Items:
   a) It is the responsibility of society, through its government, to guarantee full employment.
   b) The government should play a bigger part in the economic life of the nation in order to distribute income more equally.
   c) Poverty could almost be done away with if we made certain basic changes in our social and economic system.
   d) [Canada] needs a complete [socialist] restructuring of its basic institutions.
2. *Professional Autonomy (newly derived)*
   Internal consistency in current population: alpha = .63
   Scale: 5 = strongly agree—1 = strongly disagree
   Items (R = item is reversed for scoring purposes):
   a) The federal government has no business trying to influence the specialty choices of medical students.
   b) The best thing about medicine in the past, as compared to today, is that you could be your own boss.
   c) There are too many controls on the medical profession that interfere with taking care of patients.
   d) Government should have a strong voice in determining the organizational form of medical practice. (R)
3. *Physician Activism*[17] *(modified for this study)*
   Internal consistency in current population: alpha = .83
   Scale: 5 = definitely proper—1 = definitely improper
   Items:
   a) Placing activist pamphlets or posters in waiting rooms.
   b) Wearing buttons with activist slogans while on duty.
   c) Participating as doctors in marches or other demonstrations.
   d) Participating in strikes against individual hospitals or clinics.
   e) Participating in strikes against the provincial or federal government.
4. *Medical Altruism (newly derived)*
   Internal consistency in current population: alpha = .53
   Scale: 5 = strongly agree—1 = strongly disagree
   Items (R = item is reversed for scoring purposes):
   a) Medical care is a right, not a privilege.
   b) A physician is morally obligated to treat people of all races, religions, sexual orientations, and personal habits, regardless of his or her own preferences.
   c) It is the responsibility of society, through its government, to provide everyone with best available medical care, whether he can afford it or not.
   d) Things work best when people concern themselves mainly with their own welfare and let others take responsibility for themselves. (R)
5. *Social Benefits of Extra-Billing (newly derived)*
   Internal consistency in current population: alpha = .76
   Scale: 5 = strongly agree—1 = strongly disagree
   Items (R = item is recorded for scoring purposes):
   a) Extra-billing never presented a significant barrier to poor people seeking medical care.
   b) With extra-billing abolished, the provincial legislature is already at work seeking ways to further restrict the independence of physicians.
   c) The elderly and indigent are *more* likely to receive needed medical care now that extra-billing is prohibited. (R)
   d) The ban on extra-billing will inevitably lead to more complete government control of medicine.
6. *Impression of How Others Viewed the Strike (newly derived)*
   Internal consistency in current population: alpha = .82
   Scale: 3 = in favor of, 2 = neutral towards, 1 = opposed to
   Items:
   a) Physician partner or closest medical colleague.
   b) Spouse.
   c) Office staff.
   d) Nurses with whom I work (on average).
   e) My patients (on average).
   f) The general public (on average).
7. *Physician Satisfaction*[12] *(last 3 items newly derived)*
   Internal consistency in current population: alpha = .83
   Scale: 5 = extremely satisfied—1 = very dissatisfied
   Items:
   a) Your contact with other physicians.
   b) Your relationships with non-physician health professionals.
   c) Degree to which your work is educationally stimulating.
   d) Manpower resources available to you.
   e) Diversity of patients you see (age, sex, types of problems/disease).
   f) Patient volume.
   g) Your ability to derive personal gratification from patient care.
   h) Salary/income.
   i) Your role in making administrative decisions.
   j) Your ability to meet the needs and demands of patients.
   k) The degree of status and prestige associated with your work.
   l) Your ability to remain knowledgeable and current.
   m) Opportunities for advancement and future success.
   n) Your ability to provide patients with care of the highest quality.
   o) Your ability to hospitalize patients when you consider it necessary.
   p) Your ability to obtain clinically indicated diagnostic tests on your patients.

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ANA Opts for New Structure and Broader Membership Plan

The American Nurses' Association (ANA) House of Delegates, at its recent annual meeting, adopted a flexible new plan to alter the structure of the nation's largest professional nursing organization and its 53 constituent state nurses associations. New by laws adopted by the house will broaden the ANA membership and alter its governance structure and relationships with other national organizations.

Most of the changes come from recommendations made by the ANA Commission on Organizational Assessment and Renewal (COAR), following 18 months of study, proposal development, and consensus building by the commission with ANA's constituency, its elected and appointed officials, and other national nursing groups and the profession at large. Some of the changes wrought by the new by laws include:

- Multi-state nurses associations and organizations of US nurses overseas can become constituent members of ANA whereas only state nurses associations were eligible in the past.
- State nurses associations can have organizational affiliates or members. These may include specialty nursing organizations meeting the qualifications set forth in the by laws. The rights of these affiliates will be prescribed by the individual SNAs, and be limited to the state level of the association.
- Individual members of SNAs must be registered nurses. A proviso requires the ANA House of Delegates to determine the membership rights, responsibilities, and privileges of these individuals at the time SNAs move to admit technical nurses as members.
- At the close of next year's House of Delegates meeting, a Congress of Nursing Practice and a Congress on Nursing Economics will be formed—to focus on long-range policy development, standard setting, program development, and evaluation in their area of expertise, as well as addressing emerging health care and nursing issues on an ad hoc basis.

Other matters addressed by the 1989 ANA House of Delegates included: seating of ANA councils and organizational affiliates in the House of Delegates; establishing a separately incorporated center to serve the ANA credentialing programs; retaining authority for setting standards related to nursing education, practice and service; expansion of joint ventures with other nursing organizations; enhancing non-dues revenue and developing incentive programs to build membership; provision for increases in dues assessment paid to ANA; and receiving a progress report on plans and activities to relocate the association's headquarters to Washington, DC.

The ANA is the national professional association representing the nation's two million registered nurses through its 53 constituent state and territorial nurses associations. The 1990 ANA House of Delegates will meet June 15–20, 1990 during the association's biennial convention in Boston.

Induced Abortion and the Risk of Subsequent Ectopic Pregnancy

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Abstract: This study assessed the effect of legal induced abortion on ectopic pregnancy risk by using a comparison group of reproductive-age women who were at risk of becoming pregnant during the same time period the women with ectopic pregnancy conceived. Cases were members of Group Health Cooperative of Puget Sound who were hospitalized for ectopic pregnancy from October 1981 through September 1986 (N = 211). Controls were randomly selected members matched to cases on age and county of residence (N = 457). All subjects in this analysis had had one or more prior pregnancies.

Eighty-eight cases (41.7 per cent) and 177 controls (38.7 per cent) had a history of one or more induced abortions. The relative risk of ectopic pregnancy associated with one abortion was 0.9 (95 per cent confidence interval 0.6, 1.3), adjusted for age, county, reference date, religion, gravidity, age at first pregnancy, lifetime number of sexual partners, and miscarriage history. Among women with two or more prior pregnancies, the risk associated with two or more abortions was 1.2 (0.6, 2.4). Controlling for pelvic inflammatory disease and use of intrauterine devices did not alter these risks. We conclude that legal abortion as performed in the US since 1970 has little or no influence on a woman's risk of ectopic pregnancy in subsequent pregnancies. (*Am J Public Health* 1989; 79:1234–1238.)

Introduction

The rate of ectopic pregnancy in the United States tripled between 1970 and 1983, from 4.2 to 12.6 per 10,000 women between 15 and 44 years of age.[1,2] Although improved diagnostic capabilities may be responsible for enhanced case finding, it seems likely that this increase in rates reflects a substantial actual increase in ectopic pregnancy incidence.[3] During the same time period, the incidence of induced abortions rose from 5 per 1,000 women 15 to 44 years in 1970 to 23 per 1,000 in 1983.[4,5] This parallel increase suggests the possibility that induced abortion may predispose to a subsequent ectopic pregnancy. A possible basis for such a predisposition could be fallopian tube scar tissue formation associated with salpingitis, since induced abortion has the potential to introduce salpingitis-causing organisms via contaminated instruments or cause organisms already present in the vagina to ascend.[3,6]

The results of recent US research concerning the association between induced abortion and ectopic pregnancy have been generally reassuring, but not conclusive. Chung's prospective cohort study in Hawaii found no increased risk.[7] In case-control studies, however, both Daling, *et al*, and Levin, *et al*, found a weak association between a history of multiple abortions and ectopic pregnancy, but the studies were small and chance effects plausible.[8,9] Both case-control studies compared cases to controls chosen from women delivering a liveborn child. Such a control group may underrepresent women with a history of induced abortion and create an artifactual relation between abortion and ectopic pregnancy in the absence of a real association.[10] The present study sought to eliminate this potential bias by the use of randomly selected reproductive-age women as a comparison group.

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Methods

Case Identification

Eligible cases were all women members 18 years of age or older who were hospitalized with the diagnosis of ectopic pregnancy at Group Health Cooperative of Puget Sound hospitals between October 1, 1981 and September 30, 1986. Group Health Cooperative is a prepaid health care cooperative founded in 1947; it serves over 320,000 residents of western Washington State. Cases were identified using a computerized hospital discharge system that has been in use at Group Health Cooperative since 1972.

Each patient's primary care physician at Group Health Cooperative was sent a letter describing the study and asking permission to contact the patient. If he or she consented, the Group Health Cooperative Research Department mailed a letter with a self-addressed card to the subject, asking her to return the card to indicate whether or not she wished to participate. If she agreed to participate, the interviewer telephoned her to explain what her participation in the study would involve, answer any questions about the study, and arrange for the interview. Subjects signed consent forms and medical records releases at the interview. Of the 396 patients identified as eligible from hospital records, primary care physicians refused permission to contact 11. In addition, 39 patients refused to participate, and nine could not be reached. A total of 337 cases were interviewed, 85.1 per cent of those eligible (Table 1).

An estimate of the date a case conceived her ectopic pregnancy was considered her "reference date." All questionnaire information gathered pertained to exposures and characteristics prior to that date.

Control Identification

Controls were women of child-bearing age identified

TABLE 1—Subject Response by Case Status

Total Subjects Identified	Cases		Controls	
	N = 396	(%)	N = 1139	(%)
Interviewed	337	(85.1)	893	(78.4)
MD refusal	11	(2.8)	32	(2.8)
Patient refusal	39	(9.8)	168	(14.7)
Lost to follow-up	9	(2.3)	46	(4.0)

from Group Health Cooperative membership files. They were selected randomly, and matched to cases by age and county of residence. Ages were matched as closely as possible: within one year for the majority of subjects, extending to a maximum of five years for one set. Two potential controls were initially selected for each case, and if either of them was unreachable or declined to participate, further controls were selected. Efforts continued to enlist the participation of everyone who was initially selected. Consequently, in many instances more than two controls were matched to each case. Controls were assigned a reference date which was the midpoint of the three-month quarter containing the reference date of the case to whom they were matched (Table 2). To be eligible, a control had to be a Group Health Cooperative member at her assigned reference date. After a potential control was identified, her primary physician was notified. If the physician gave permission for her to be contacted, the same sequence of events was followed as for cases. Of 1,139 potential controls identified, 32 were not contacted due to physician request. When contacted, 168 women refused to participate, and 46 could not be reached. Eight hundred ninety-three control subjects were interviewed, 78.4 per cent of those eligible (Table 1).

Potential controls whose husbands or partners were sterilized were excluded from consideration if the sterilization had not been reversed. Women with a history of hysterectomy or no sexual activity for the 12 months preceding the reference date were also excluded from the control group. Also, women four to nine months pregnant at the reference date, and therefore not at risk of being diagnosed as having an ectopic pregnancy, were identified from Group Health Cooperative prenatal files and excluded as potential control subjects.

Data Collection

Ninety-nine per cent of the cases and 95 per cent of the controls were interviewed in person. During the interviews, which lasted an average of 56 minutes, respondents were queried concerning their reproductive, contraceptive, sexual, and medical histories, past hygienic practices including douching, and demographic and lifestyle characteristics. For each induced abortion, information obtained included gestational age, abortion method, medical setting where the procedure was performed, and whether infection occurred after the abortion.

To decrease the possibility of recall bias, month by month calendars of all contraceptive, reproductive and other important personal events were formulated and discussed by interviewers and all subjects. The interviewers were not blinded to the subjects' diagnoses, but when 24 cases and 66 controls were contacted again and reinterviewed with selected questions by a different interviewer, the responses were in agreement 97 per cent of the time for cases and 95 per cent for controls.

As part of a substudy on the validity of self-reported data, the medical records of all subjects who reported intrauterine device (IUD) use and a random sample of 20 per cent of other subjects were reviewed. Information obtained from medical charts included history of induced abortion. Interview and chart information agreed for 264 of 285 subjects for whom there was information from the chart (Table 3).

Analysis

For this analysis, cases (n = 23) and controls (n = 14) with a prior ectopic pregnancy were excluded. Also excluded from both groups were women without a pregnancy prior to reference date (65 cases and 263 controls), and women who had had an induced abortion before 1970 when abortion became legal in Washington State (two cases and 15 controls). All surgically sterilized women (35 cases and 153 controls) were excluded since they already had their fallopian tubes occluded, the hypothesized means by which induced abortion could predispose to an increased risk of ectopic pregnancy. Seven cases with nontubal ectopic pregnancies (six ovarian and one abdominal) were excluded. Five subjects (one case and four controls) were excluded due to missing data for key model variables. Two hundred ten cases and 453 controls remained for analysis.

Multiple logistic regression analysis was performed to control for matching variables (age, reference date, county) and to adjust for the confounding effect of other exposures.[11] Relative risks approximated by odds ratios were estimated by the method of maximum likelihood. Significance tests were based on the chi-squared approximation for the likelihood ratio statistic, and 95 per cent confidence intervals were based on the standard error of coefficient estimates and the normal approximation.

Dichotomous variables considered in the analysis were education (≤12 years vs >12 years), marital status at refer-

TABLE 2—Case and Control Comparability on Matching Variables

Variables	Cases N = 211	Cases (%)	Controls N = 457	Controls (%)
Age (years) at reference date				
<20	7	(3.3)	3	(0.7)
20–24	34	(16.1)	56	(12.3)
25–29	55	(26.1)	133	(29.1)
30–34	83	(39.3)	183	(40.0)
35+	32	(15.2)	82	(17.9)
Reference year				
1981	16	(7.6)	19	(4.2)
1982	50	(23.7)	98	(21.4)
1983	36	(17.1)	92	(20.1)
1984	50	(23.7)	103	(22.5)
1985	41	(19.4)	91	(19.9)
1986	18	(8.5)	54	(11.8)
County of residence				
King	140	(66.4)	307	(67.2)
Pierce	28	(13.3)	71	(15.5)
Snohomish	19	(9.0)	27	(5.9)
Thurston	20	(9.5)	45	(9.8)
Other	4	(1.8)	7	(1.5)

TABLE 3—Comparison of Induced Abortion History Between Interview and Medical Chart Information

	Medical Chart: No[a]		Medical Chart: Yes		Medical Chart: Unknown	
Interview	N	(%)	N	(%)	N	(%)
Cases						
No	66	(57.9)	5	(4.4)	3	(2.6)
Yes	3	(2.6)	35	(30.7)	2	(1.8)
Controls						
No	106	(59.9)	4	(2.3)	0	
Yes	9	(5.1)	57	(32.2)	1	(0.6)

[a]Includes women whose medical record data does not cover the same pre-reference date time period reported in the interview.

ence date (married, living as married or steady partner vs separated, divorced, or never married), ever-use of oral contraceptives, ever-use of Dalkon Shield or other IUD, current use of contraceptives, history of douching, miscarriage, gonorrhea, pelvic inflammatory disease (PID), and endometriosis. Two dichotomous variables were used for abortion history: one abortion vs none, and two or more vs none. Categorical variables included cigarette smoking (never, former, current), race (Black, White, other), religion (Catholic, Protestant, other, none), age (<20 years, 20–24, 25–29, 30–34, 35+), family income (<$15,000 per year, $15–30,000, >$30,000), reference year (1981–82, 1983–84, 1985–86), duration of barrier contraceptive use (no use, <2 years, 2–4 years, 5+ years), lifetime number of sex partners (one, 2–4, 5+), age at first regular intercourse (<18 years, 18–20, 21+), age at first pregnancy (<20 years, 20–24, 25+), and gravidity (one, two, three or more pregnancies).

Results

A higher percentage of cases than controls were Black (Table 4). Cases were also more likely to have a family income under $15,000 per year and to be current cigarette smokers. Average age at first regular intercourse and age at first pregnancy were lower for cases than controls, while the cases had had a higher lifetime number of sexual partners. Case women also reported a history of PID more frequently than control women, with 4.7 per cent of cases and 3.3 per cent of controls reporting a first episode of PID occurring after their first abortion. IUD ever-use was slightly greater among cases, due to their higher frequency of Dalkon Shield IUD use. Controlling for either PID history or IUD usage did not alter the risk of ectopic pregnancy associated with induced abortion. Cases were more than three times as likely not to be contracepting at reference date than controls (Table 4), but abortion history did not differ between women using and those not using contraception at reference date (Table 5).

Eighty-eight cases (41.7 per cent) and 177 controls (38.7 per cent) had had one or more induced abortions (Table 6). The relative risk of ectopic pregnancy associated with a history of one induced abortion was 0.9 (95 per cent confidence interval 0.6, 1.3) when adjusted for age, county of residence, reference date, religion, age at first pregnancy, number of pregnancies, lifetime number of sexual partners, and prior miscarriage history. Among women with two or more prior pregnancies (125 cases and 288 controls), the relative risk for two or more abortions was 1.2 (95 per cent CI = 0.6, 2.4).

Method of abortion, gestational length, and site where abortion was performed did not vary appreciably between cases and controls (Table 7). Cases reported pain for longer than one week after last abortion twice as often as controls, but this result could be due to chance. The mean time from last abortion to reference date was 64.8 months for cases and 68.8 months for controls.

Discussion

We found little if any increased risk of ectopic pregnancy associated with prior induced abortion. Generalization of these results to a population outside the United States or to one for whom induced abortion is not legal may be misleading since the manner in which abortions are performed may influence their effect on ectopic pregnancy risk.

Ascertainment of abortion exposure was based on subject self-report in this study. If selective underreporting by cases or controls existed, our estimates of relative risk may be biased. Information from the substudy on validity of self-reported data indicates a small amount of underreporting

TABLE 4—Selected Demographic Characteristics

Demographic Characteristics	Cases N = 211	Cases (%)	Controls N = 457	Controls (%)
Race				
White	171	(81.0)	392	(85.8)
Black	29	(13.7)	31	(6.8)
Other (Asian, Hispanic, Native American, others)	11	(5.2)	34	(7.4)
Religion[a]				
Catholic	45	(21.3)	87	(19.1)
Protestant	88	(41.7)	225	(49.5)
None	63	(29.9)	106	(23.3)
Other	15	(7.1)	37	(8.1)
Education				
≤12 years	66	(31.3)	135	(29.5)
>12 years	145	(68.7)	322	(70.5)
Family Income[b]				
<$15,000/year	48	(22.7)	70	(15.4)
$15–30,000	92	(43.6)	211	(46.4)
>$30,000	71	(33.6)	173	(38.0)
Cigarette Smoking Status				
Never	89	(42.2)	238	(52.1)
Former	41	(19.4)	89	(19.5)
Current	81	(38.4)	130	(28.4)
Marital Status				
Married	142	(67.3)	321	(70.2)
Living as married	27	(12.8)	34	(7.4)
Separated	5	(2.4)	9	(2.0)
Divorced	7	(3.3)	23	(5.0)
Never married	5	(2.4)	24	(5.3)
Steady partner, living separately	25	(11.8)	46	(10.1)
History of PID[c]	33	(15.6)	37	(8.1)
IUD ever-use[d]				
None	134	(63.5)	311	(68.1)
Dalkon Shield	25	(11.8)	26	(5.7)
Non-Dalkon Shield only	50	(23.7)	117	(25.6)
Number of pregnancies before reference date				
1	86	(40.8)	169	(37.0)
2	60	(28.4)	150	(32.8)
3	31	(14.7)	83	(18.2)
4 or more	34	(16.1)	55	(12.0)
Number of miscarriages before reference date				
0	142	(67.3)	374	(81.8)
1	50	(23.7)	63	(13.8)
2 or more	19	(9.0)	20	(4.3)
Number of births before reference date				
0	60	(28.4)	103	(22.5)
1	91	(43.1)	165	(36.1)
2	43	(20.4)	133	(29.1)
3 or more	17	(8.1)	56	(12.3)
Age at first pregnancy				
<21 years	102	(48.3)	151	(33.0)
21–24 years	56	(26.5)	171	(37.4)
25+ years	53	(25.1)	135	(29.5)
Age at first regular intercourse				
<18 years	63	(29.9)	103	(22.5)
18–20 years	99	(46.9)	203	(44.4)
21+ years	49	(23.2)	151	(33.0)
Lifetime number of sexual partners[e]				
One	27	(12.8)	125	(27.4)
Two to four	75	(35.5)	146	(31.9)
Five or more	108	(51.2)	184	(40.3)
Contracepting at reference date				
Yes	44	(20.9)	320	(70.0)
No	167	(79.1)	137	(30.0)

[a]Two controls were uncertain. [b]Three controls refused to answer this question. [c]One case and six controls were uncertain. [d]Two cases and three controls were uncertain of type used. [e]One case and two controls refused to answer this question.

TABLE 5—Abortion History by Contraceptive Use at Reference Date

Abortion History	Not contracepting at reference date N = 304	(%)	Contracepting at reference date N = 364	(%)
Number of Induced Abortions				
0	186	(61.1)	217	(59.6)
1	82	(27.0)	100	(27.5)
2 or more	36	(11.8)	47	(12.9)

TABLE 6—Standardized Relative Risks of Ectopic Pregnancy According to Abortion History

	Cases N = 211	(%)	Controls N = 457	(%)	Relative Risk* (95% CI)
Number of Induced Abortions					
0	123	(58.3)	280	(61.3)	1.0
1	55	(26.1)	127	(27.8)	0.9 (0.6, 1.3)
2 or more	33	(15.6)	50	(10.9)	1.2** (0.6, 2.4)

*Adjusted for age, county, reference date, religion, gravidity, age at first pregnancy, lifetime number of sexual partners, and prior miscarriage history, using a multiple logistic regression model.

**Compared to women with two or more prior pregnancies, no induced abortions.

TABLE 7—Characteristics of Last Induced Abortion in Women with One or More Abortions

Characteristics	Cases N = 88	(%)	Controls N = 177	(%)
Abortion Method				
Suction	71	(80.7)	137	(77.4)
Dilatation and Curettage	9	(10.2)	19	(10.7)
Saline Infusion	5	(5.7)	8	(4.5)
Prostaglandin Infusion	1	(1.1)	1	(0.6)
Other	1	(1.1)	1	(0.6)
Uncertain	1	(1.1)	11	(6.2)
Complications				
Infection	2	(2.3)	9	(5.1)
Bleeding > one week	2	(2.3)	6	(3.4)
Pain > one week	11	(12.5)	11	(6.2)
D & C after abortion	3	(3.4)	7	(4.0)
Hospitalization after abortion	0	(0.0)	4	(2.3)
Length of Gestation				
First trimester	81	(92.0)	165	(93.2)
Second trimester	7	(8.0)	12	(6.8)
Median length (wks)	8.0		8.0	
Laminaria Used	20	(22.7)	35	(19.8)
Site abortion performed				
MD office	14	(15.9)	44	(24.9)
Freestanding clinic	37	(42.0)	65	(36.7)
Hospital or hospital clinic	36	(40.9)	68	(38.4)
Mean time from last abortion to reference date (months)		64.8		68.8

of prior induced abortion status may have been present. Four per cent of cases and 2 per cent of controls denied a history of induced abortion during their study interview, but had such a history noted in their medical charts. If the chart information more closely represents actual abortion history, our results may be a slight underestimate of the true effect. In our study, 42 per cent of cases and 39 per cent of controls stated they had had one or more induced abortions, whereas a study of Washington state multigravida women giving birth in 1984 and 1985 found that 25.6 per cent reported a prior abortion on the birth certificates in a random sample.[12] Since our subjects reported a considerably higher frequency of abortion history, we may be somewhat reassured that substantial underreporting does not exist.

Recent US studies have disagreed about the extent to which abortion is a risk factor for subsequent ectopic pregnancy. In case-control studies, Daling, *et al*, and Levin, *et al*, both found a modestly elevated ectopic pregnancy risk in women with a history of one abortion (Daling RR = 1.4, Levin RR = 1.3), and a somewhat higher risk associated with multiple induced abortions (Daling RR = 1.8, Levin RR = 2.6).[8,9] However, in both studies the number of cases with two or more abortions was small and associations found could have been attributable to chance. Moreover, both researchers chose controls delivering a child at a time comparable to the case's ectopic pregnancy surgery, and by definition this control group did not include women who conceived but chose to terminate that pregnancy. The cases had no similar termination opportunity, as ectopic pregnancy usually manifests itself before a woman's knowledge of her pregnant status. Since women terminating pregnancies have been estimated to have a higher frequency of past abortions than do women carrying a full-term pregnancy, use of these control groups may have led to an artificially elevated ectopic pregnancy risk associated with prior abortion.[10] Although Daling, *et al*, attempted to correct for this bias by limiting their study to married women and controlling for contraceptive use at time of conception, we believe our use of a more general population control group is a more direct way to address this issue.

Chung's Hawaiian study followed a large cohort of women who had abortions performed in the four years immediately following legalization of the procedure in Hawaii in 1970. The researchers found little increased risk associated with abortion in general (RR = 1.2), but did find a strong association (RR = 5.1) between abortion and ectopic pregnancy in women who had post-abortion infection or retained products of conception.[7] In our study, post-abortion infection was identified by 2.3 per cent of the cases and 5.1 per cent of the controls who had a history of abortion, and was not a risk factor for subsequent ectopic pregnancy. One difference between our study and Chung's which may account for this discrepancy is the time period during which the study was conducted. As years pass since the legalization of induced abortion in Washington State in 1970, physician expertise in the performance of the procedures may be improving, thereby minimizing risk to the patient. The abortion procedure itself has also changed over the past two decades, with decreased use of dilatation and curettage (used in 10 per cent of our cases' and 11 per cent of our controls' last abortion) and increased use of less traumatic suction procedures (used in 81 per cent of cases' and 77 per cent of controls' last abortion).

In conclusion, this study yields no evidence that legal abortions performed since 1970 in our study population led to increased risk of ectopic pregnancy in subsequent pregnancies. We found no increase in risk in women with a history of one induced abortion (RR = 0.9), and the slightly increased risk we found to be associated with multiple abortions (RR = 1.2) could easily have been due to chance. Providing that the

Group Health Cooperative experience is similar to that of other populations within the US, these results argue that the increasing rate of induced abortions is not an important factor in explaining the sharp rise in ectopic pregnancy incidence in this country in the past two decades.

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Yale University to Strengthen Its Public Health Program

A Yale University committee released a report in mid-July recommending to President Benno C. Schmidt Jr. a strategy to strengthen teaching and research in public health at the institution. The committee, which deliberated over six months, gathered information from faculty, students, alumni, and local and national leaders in public health. The basic goals of the committee's recommendations are to establish a leadership role for Yale in the arena of public health. In its report, the committee emphasized that the public health profession plays a critically important role in protecting and improving the physical and mental well-being of the nation.

Specifically, the committee recommended that the Department of Epidemiology and Public Health remain a department of the School of Medicine within which it will retain its status as a school for purposes of accreditation. To signal clearly the comprehensive mission of the department, the Committee suggests that the department be renamed the Department of Public Health. Burton H. Singer, PhD, professor of public health and the head of the biostatistics division, was named to a two-year term as Chairman of the Department of Public Health and Associate Dean for Public Health. A native of Chicago, Dr. Singer received BS and MS degrees from Case Institute of Technology in 1959 and 1961, respectively. In 1967, he received a PhD degree in statistics from Stanford University; he joined the Yale faculty in 1985. Dr. Singer succeeds Jan A. J. Stolwijk, PhD, who has completed seven years as Chairman and previously announced that he would return to full-time teaching and research.

Other recommendations for public health at Yale that the president will consider include:

- Establishing a standard committee within the School of Medicine to ensure ways of integrating teaching and research programs in public health with medicine;
- Allocating new resources to the department for the purpose of renovation of facilities and the recruitment of faculty;
- Strengthening the process for recruitment and admission of students;
- Strengthening ties between the Department of Public Health and other interested units of the University;
- Renovation and expansion of public health facilities at 60 College Street for public health programs;
- Increasing the size of the faculty by recruiting up to eight new faculty members;
- Strengthening the Master of Public Health Program by reorganizing the student admission process and focusing on the development of leaders in the field.

Founded in 1915, the present Department of Epidemiology and Public Health is a nationally accredited school of public health. It is among the oldest of 23 such schools in the United States and one of the few that is a department within a medical school.

The Relationship of Smoking and Ectopic Pregnancy

ARDEN HANDLER, DRPH, FAITH DAVIS, PHD, CYNTHIA FERRE, MA, AND TIMOTHY YEKO, MD

Abstract: A case-control study, using data abstracted between 1983 and 1987 from a large perinatal registry, was conducted to explore the relationship between smoking and ectopic pregnancy. Women with ectopic pregnancy (n = 634) seen at University of Illinois Perinatal Network Hospitals were compared to women who were delivered of a single live-born infant (n = 4287). Adjusted for age and race, women who reported smoking during pregnancy had a greater than twofold risk of ectopic pregnancy (Odds Ratio = 2.5, 95% confidence interval = 1.9, 3.2) compared to women who never smoked. The estimated relative risk rose from 1.4 (95% CI = 0.8, 2.5) for a woman smoking fewer than 10 cigarettes a day to 5.0 (95% CI = 2.9, 8.7) at one and a half or more packs of cigarettes per day (p-value for trend <0.001). Although further basic and epidemiologic research is necessary, the observed dose-response relation strengthens the argument that smoking may be a causal factor in the development of ectopic pregnancy. (*Am J Public Health* 1989; 79:1239–1242.)

Introduction

Although ectopic pregnancy is a relatively uncommon occurrence, its incidence has quadrupled from a rate of 5/1000 live births in 1970.[1] Simultaneously, maternal mortality rates from other causes have fallen, making ectopic pregnancy the primary cause of maternal mortality during the first trimester of pregnancy. Known risk factors for ectopic pregnancy include pelvic inflammatory disease, prior tubal surgery, contraceptive history, and a history of infertility.[2] As the rise in incidence of ectopic pregnancy has paralleled the increase in cigarette smoking among women, it is reasonable to examine smoking as a potential risk factor.

Most of the evidence for the positive association of smoking and ectopic pregnancy derives from studies in which smoking was not a primary hypothesis (Odds Ratios = 1.5 to 3.9).[3–6] Recently Chow, *et al*,[7] investigated the potential etiologic relationship between cigarette smoking at the time of conception and ectopic pregnancy, considering the effect of age at initiation, smoking duration, and intensity on this event. An overall association of smoking with ectopic pregnancy was evident (OR = 2.2), but a dose-response was not demonstrated, possibly due to an inadequate number of cases. To permit more stable estimates of the risk of ectopic pregnancy by level of smoking, a case-control study using data from a large perinatal registry was conducted.

Methods

The University of Illinois Hospital (UIH) Perinatal Network collects information on 425 variables for all maternal and neonatal events occurring at its 12 member hospitals. Data are abstracted daily from medical records by trained abstracters and entered into the UIH perinatal database.

Women with an ectopic pregnancy (n = 701) presenting at UIH Perinatal Network Hospitals from 1983–87 were identified. Cases with a prior ectopic pregnancy were excluded (n = 67), leaving 634 total cases for study. Controls were a 5 percent random sample of all women without a prior history of ectopic pregnancy delivered of a single live-born infant (n = 4287) at these same hospitals. Women with pregnancies resulting in spontaneous abortions or fetal deaths were excluded for use as controls in this study because of the known relationship between maternal smoking and these reproductive outcomes.[8,9]

Cigarette smoking data in the UIH perinatal database is a measure of smoking during the current pregnancy. It is recorded as a dichotomous variable (yes/no) and as a continuous variable (number of cigarettes smoked per day). Ascertainment of smoking history was similar between cases and controls. Information on smoking as a dichotomous variable was available for 93.4 percent of the cases and for 91.9 percent of the controls; information on cigarette dose was available for 87.9 percent of the cases and 87.6 percent of the controls. Five levels of cigarette smoking were considered: nonsmokers; less than 10 cigarettes/day; 10–19/day; 20–29/day; and 30 or more/day.

Other variables collected on the perinatal abstract form were examined as potential risk factors: maternal age, race, payment method (as an indicator of socioeconomic status), gravidity, parity, prior spontaneous abortion, and prior induced abortion. Past or present pelvic inflammatory disease, past and current contraceptive use, and detailed fertility history were not available for this analysis.

The odds ratio and 95% Taylor series confidence intervals were used to describe the association between ectopic pregnancy and smoking.[10,11] Adjusted odds ratios were calculated using the Mantel-Haenszel technique and 95% confidence intervals according to the method of Robins, Greenland, and Breslow.[12] Tests for trend were computed by the Mantel extension.[13] To adjust simultaneously for multiple factors, a multiple logistic regression model was fitted with smoking as a dichotomous independent variable (yes/no) and ectopic pregnancy as the dependent variable.[14] The model was restricted to women with no prior history of spontaneous abortion since spontaneous abortion history modified the smoking-ectopic relationship.

An adjusted dose-response relationship between smoking and ectopic pregnancy was tested in a second multivariable model, with cigarette smoking treated as a continuous variable. In a third multivariable model, smoking was treated as a series of dummy variables to estimate dose-specific odds ratios.

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Results

The rate of ectopic pregnancy at the University of Illinois Perinatal Network hospitals from 1983 to 1987 was 7.9 per 1000 live births. Increasing maternal age, Black race, increased gravidity, increased parity and prior spontaneous abortion are risk factors for ectopic pregnancy in this population (Table 1). Asians had a lower risk of ectopic pregnancy than other ethnic groups.

Women who reported smoking during pregnancy had a greater than twofold risk of ectopic pregnancy. Controlling for maternal age, race, payment method, gravidity, parity, and spontaneous abortion separately did not alter this crude estimate (range of estimates = 2.2–2.4). Smoking raised the risk of ectopic pregnancy less among women who had a prior spontaneous abortion (OR = 1.6, 95% CI = 1.1, 2.3) than among women who had not (OR = 2.5, 95% CI = 2.1, 3.1). This is due to the higher prevalence of smoking among women with ectopic pregnancy who have *not* had a prior spontaneous abortion (Table 1).

The smoking effect among women with no prior history of spontaneous abortion, adjusting for maternal age and race (Black versus other) was OR 2.5 (95% CI = 1.9, 3.2) (Table 2). A dose-gradient between cigarette smoking and ectopic pregnancy was evident when smoking was evaluated as a continuous variable (regression coefficient = 0.049, S.E. = 0.006) or as a categorical variable (Table 2 and Figure 1).

Discussion

In this study, women who reported smoking during pregnancy had a greater than twofold risk of ectopic pregnancy. This effect is similar to that seen in previous studies (Table 3).[4–7] The reported association is strengthened by the dose-response after adjusting for maternal age and race. Based on the estimates presented here, 60 percent of the ectopic pregnancies among smokers may be due to smoking.[11]

The significantly increased risk of ectopic pregnancy associated with increasing age and with Black race demonstrated in this report is consistent with other studies of ectopic pregnancy.[2] The gravidity and parity effect observed in the crude analysis disappeared when adjusting for prior spontaneous abortion. While in this population there was no increased risk of ectopic pregnancy associated with prior induced abortion, an increased risk with prior spontaneous abortion among multigravid women was observed. Previous

TABLE 1—Distribution of Cases and Controls by Selected Risk Factors and Smoking Status, Odds Ratios (OR), and 95% Confidence Intervals for Each Factor and Ectopic Pregnancy

	Cases (N = 634)		Controls (N = 4287)			
Risk Factors	n	% smoke[a]	n	% smoke[a]	OR	95% CI
Maternal age[b] (years)						
24 or Less	166	47	1618	27	1.0	
25–34	369	42	2268	22	1.6	1.3, 1.9
Over 35	99	32	400	19	2.4	1.8, 3.2
Missing	0		1			
Race						
White	317	47	2171	27	1.0	
Black	196	45	974	31	1.4	1.1, 1.7
Hispanic	103	26	849	13	0.8	0.7, 1.1
Asian	17	0	205	4	0.6	0.3, 0.9
Other/Missing	1		88			
Payment Method						
Private	349	38	2156	19	1.0	
Medicaid	140	52	1074	35	0.8	0.7, 1.0
Self	98	44	548	23	1.1	0.9, 1.4
Other	47	30	422	21	0.7	0.5, 0.9
Missing	0		87			
Gravidity[b]						
Primigravida	139	38	1298	19	1.0	
1 Prior	168	45	1237	21	1.3	1.0, 1.6
2 or More Prior	318	41	1667	29	1.8	1.4, 2.2
Missing	9		84			
Parity[b]						
Nulliparous	229	43	1730	22	1.0	
Primipara	188	41	1305	23	1.1	0.9, 1.3
2 or More Prior	208	41	1170	27	1.3	1.1, 1.6
Missing	9		82			
Prior Spontaneous Abortion[c]						
Yes	163	37	749	28	1.5	1.2, 1.8
No	323	45	2158	25	1.0	
Prior Induced Abortion[c]						
Yes	76	58	515	36	0.9	0.7, 1.1
No	410	40	2392	24	1.0	
Current Smoking						
Yes	246	39	932	22	2.3	1.9, 2.7
No	346		3008		1.0	
Missing	42		347			

[a]% Smokers in each category for cases and controls.
[b]Mantel Extension $p < 0.001$.
[c]For multigravid women only.

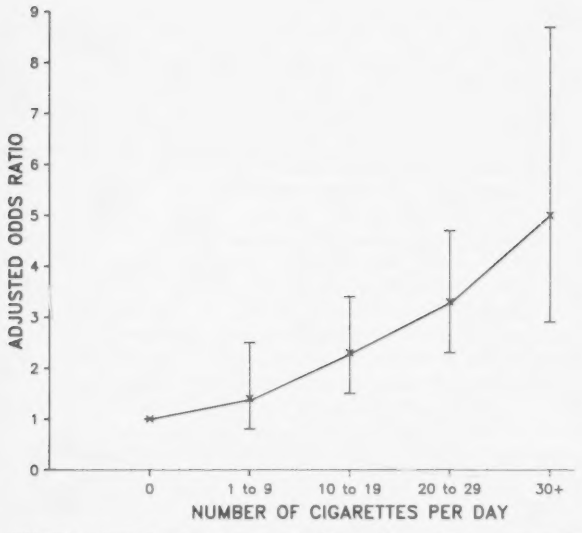

FIGURE 1—Dose Response Curve for Smoking and Ectopic Pregnancy among Women with no Prior Spontaneous Abortion, Adjusted for Maternal Age and Race.

TABLE 2—Frequency of Cases and Controls, Adjusted Odds Ratios (OR_{adj}), and 95% Confidence Intervals for Smoking and Ectopic Pregnancy among Women with no Prior Spontaneous Abortion

Smoking Status	Cases (N = 323) n	Controls (N = 2158) n	OR_{adj}[a]	95% CI
Current Smoking (Yes/No)	152	645	2.5	1.9, 3.2
Number of Cigarettes/Day[b]				
Nonsmoker	171	1513	1.0	
<10	16	94	1.4	0.8, 2.5
10–19	33	135	2.3	1.5, 3.4
20–29	53	146	3.2	2.3, 4.7
30+	22	41	5.0	2.9, 8.7
Missing	28	229		

[a]Adjusted for maternal age and race.
[b]$p < 0.001$ for smoking as a continuous variable.

TABLE 3—Comparisons with Other Studies

Other Studies	Crude OR	95% CI
Matsunaga 1980[4]	2.3	1.0–5.5
Levin 1982[5]	1.5	0.8–2.8
WHO 1985[6]	2.1	1.7–2.6
Chow 1988[7]	2.2	1.4–3.4
Handler 1989	2.3	1.9–2.7

studies of the relationship between ectopic pregnancy and abortion of either type have been equivocal.[2,5,13–19*]

A major strength of this study is the large and ethnically diverse sample. Data collection was conducted without knowledge of the study's hypothesis, thereby preventing bias due to differential data collection for cases and controls. Despite the unexpected lower rate of ectopic pregnancy in this population, it is unlikely to affect the ascertainment of smoking data or the internal validity of the comparisons.

It is possible the exclusion of spontaneous abortions, antepartum and intrapartum fetal deaths in the control group led to an overestimate of the smoking effect since smoking increases the risk of spontaneous abortion and fetal deaths. A separate analysis including these outcomes in the control group did not alter the dose gradient observed.

The smoking data reported here refer to behavior during the current pregnancy as ascertained from medical records at the termination of pregnancy rather than interview or biochemical measures. For those identified as smoking during pregnancy, smoking behavior at conception and the duration of smoking prior to conception were unknown. Although these factors may be influential,[7] they seem unlikely to entirely explain the association with level of cigarette smoking. Additionally, women who stopped smoking just prior to pregnancy were not identifiable and their inclusion in the non-smoking group may lead to a conservative estimate of the relationship between smoking and ectopic pregnancy.

A weakness of this study is that three previously identified risk factors for ectopic pregnancy were not available in the University of Illinois Perinatal Network database and therefore could not be evaluated. These are prior or current contraceptive use, past or current history of pelvic inflammatory disease (PID), and past or current sexually transmitted diseases (STD). The World Health Organization's multinational study[6] of ectopic pregnancy found that adjustment for past history of PID or STD and for intrauterine device (IUD) use at the time of conception did not alter estimates of risk for current cigarette smokers when compared to pregnant (RR = 3.1) or nonpregnant (RR = 1.8) controls. Chow, *et al*, demonstrated an increased risk of tubal pregnancy associated with current smoking which was independent of past IUD use.

If the association between smoking and ectopic pregnancy is causal, the biologic mechanism might be one of several suggested by human and animal studies. In theory, any anatomic or physiologic factor impeding the transit of the ovum or blastocyst to the uterus may lead to an ectopic pregnancy.[20] Alterations in normal tubal architecture and tubo-ovarian relationships can mechanically delay ovum transport. This may explain the high incidence of ectopic pregnancy associated with tubal infection, pelvic surgery, and endometriosis which may cause partial tubal occlusions and/or pelvic adhesions.[21] However, up to 50 percent of the ectopic pregnancies are unexplained by any identifiable anatomical tubal factor.[22]

In humans, cigarette smoking can alter tubal motility by increasing muscular tone and the wave amplitude of tubal contractions as determined by uterotubal insufflation before and after smoking. Nicotine exposure leads to a short period of increased uterotubal activity followed by a longer period of tubal inactivity, providing increased opportunity for tubal implantation.[23]

Non-human primate and rodent studies have also demonstrated an alteration from baseline levels of tubal motility following injection of nicotine.[24–27] Nicotine treatment has been shown to delay ovum entry into the uterus, blastocyst formation, and blastocyst implantation as well as to alter spacing of implantation sites. These studies indicate cigarette smoking and nicotine are capable of affecting motility and could disrupt the complex sequence of coordinated tubal

*Holt V, Daling J, Voigt L: Induced abortion and the risk of subsequent ectopic pregnancy. Presented at the 21st Annual Meeting of the Society for Epidemiologic Research, Vancouver, Canada, June 1988.

contractions necessary for normal ovum transport.

Alternatively the effect of smoking on the occurrence of ectopic pregnancy may be due to the impairment of cellular and humoral immunity.[28–31] Smoking may be part of the causal chain linking pelvic inflammatory disease to ectopic pregnancy. It is possible that altered immunity in cigarette smokers may increase the incidence and prevalence of both tubal infection and pelvic inflammatory disease. However, the hypothesized association between smoking and pelvic inflammatory disease has not yet been explored.[7]

Basic research is needed to clarify the specific mechanism by which smoking may "cause" ectopic pregnancy. Epidemiologic studies controlling for all known risk factors are needed to confirm the dose-response relationship observed in this study between smoking and ectopic pregnancy.

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National Maternal and Child Oral Health Workshop

A national workshop entitled "Maternal and Child Oral Health" will be held September 10–12, 1989 in Washington, DC. The Workshop will solicit expert scientific and policy opinion from invited participants concerning mechanisms to optimize oral health for the maternal and child population.

The purpose of the National Workshop is to provide a set of recommendations and health policy proposals to the US Public Health Service which, when implemented, will improve the oral health of mothers and children. A report of the final recommendations will be disseminated for use by agencies, organizations, health care providers, and the public.

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Maternal Smoking during Pregnancy: No Association with Congenital Malformations in Missouri 1980–83

MICHAEL H. MALLOY, MD, MSc, JOEL C. KLEINMAN, PhD, JANICE M. BAKEWELL, WAYNE F. SCHRAMM, MA, AND GARLAND H. LAND, MPH

Abstract: Using a multisource birth defects registry developed by the Missouri Center for Health Statistics for the years 1980–83, we examined the relation between maternal smoking during pregnancy and the occurrence of congenital malformations. There were 288,067 live singleton births in this data set of which 10,223 had one or more congenital malformations. When adjusted for potential confounders the odds ratio for congenital malformations in the infants of women who smoked during pregnancy was not increased (odds ratio = 0.98, 95% confidence interval = 0.94 – 1.03). We examined the relation between smoking and groups of malformations using the International Classification of Diseases, 9th Revision, as well as analyzing for certain specific malformations within each group and found no increased risk for infants of smokers. (*Am J Public Health* 1989; 7[illegible]43–1246).

Introduction

Maternal cigarette smoking during pregnancy has been positively associated with low birthweight infants[1,2] and increased infant mortality.[3,4] The relation between smoking and the occurrence of congenital malformations is less certain. The biologic rationale for exploring the relation between smoking and congenital malformations is based on the results of experiments exposing animals to various teratogens that occur in cigarette smoke.[5–8] These teratogens include among others nicotine and cadmium which have been associated with cleft palate, anophthalmia, exencephaly, limb defects, and rib fusions.[7,8] Carbon monoxide, a major constituent of cigarette smoke has been associated with minor limb defects, but is considered more of an embryotoxic agent than a teratogen.[5,9] Clinical reports have suggested a higher prevalence of congenital heart disease in the infants of smokers[10] as well as a higher prevalence of oral clefts.[11] Other clinical reports have found no association between maternal smoking during pregnancy and congenital malformations.[12,13]

In order to examine the relation between maternal cigarette smoking and congenital malformations, we used a multisource birth defects file from Missouri. This file gives more complete ascertainment of birth defects than birth certificates alone and also contains information on maternal smoking during pregnancy.

Methods

The Missouri Birth Defects Registry for the years 1980 to 1983 is a linked file of birth certificates, hospital discharge diagnoses, neonatal intensive care unit diagnoses, Crippled Children Service diagnoses, and death certificates for children under two years of age. Linkage to the birth certificate was accomplished by matching last names and date of birth for death certificates, neonatal intensive care unit files, and Crippled Children Service files. Because names were not available on the hospital discharge file, matching was based on hospital of birth, date of birth, mother's zip code, and the child's sex and race. Match rates for the four files to the birth certificate approximated 90 percent.

This analysis was limited to singleton births. The total number of singleton births over the four-year period was 302,971. We excluded all records that were incomplete for the variables in our analysis. We were left with 288,067 records or 95 percent of the total singleton births. Variables used in this analysis and obtained from the birth certificate include race, the mother's martial status, educational attainment, age, and parity. Parity was defined as the number of live births plus abortions (induced and spontaneous) and stillbirths experienced by a woman. Information on the mother's smoking status during pregnancy was ascertained from the birth certificate as reported by the mother in terms of packs of cigarettes smoked per day. The categories recorded were nonsmoker, less than one pack per day, or one or more packs per day.

Malformations from this study were limited to those within the Version Nine - International Classification of Diseases (ICD-9) codes Congenital Malformation Chapter and to the ICD-9 Digestive System Chapter.[14] Malformations were analyzed as groups according to the 12 subchapters of the ICD-9 Congenital Malformation Chapter. These groupings attempt to categorize based on the anatomical relatedness of various specific defects (see Appendix). In addition, we analyzed several specific malformations that had been previously reported to be related to maternal smoking.

Multiple logistic regression was used to assess the effects of maternal smoking on congenital malformations, adjusting for maternal characteristics (age, marital status, educational attainment, and parity).[15] Separate models were fit for all malformations, for each group of malformations, and for each specific malformation. We ran separate models for Blacks and Whites for all malformations, groups of malformations, and specific malformations. The odds ratio for malformations for Black smokers, however, did not appear different from those of Whites.* The results presented thus combine both races with an adjustment for race within each model. Because parity had been observed to have significant interactions with maternal age in previous work done on infant mortality,[3] we fit models with two-way interactions between parity and maternal age. As in previous work, we also add a term to control for high parity (parity 3 or more for mothers under 25 years of age and parity 4 or more for those 25 or over).

Address reprint requests to Michael H. Malloy, MD, Prevention Research Program, National Institute of Child Health and Human Development, Executive Plaza North, Rm 640, Bethesda, MD 20892. Dr. Kleinman is with the Division of Analysis, National Center for Health Statistics; Ms. Bakewell, Mr. Schramm, and Mr. Land are with the Missouri Department of Health, Jefferson City. This paper, submitted to the Journal August 2, 1988, was revised and accepted for publication April 5, 1989.

*Data available on request to authors.

Results

There were 10,223 infants (3.5 percent of live births) with one or more congenital malformations ascertained from the multisource file and 13,582 unique malformations (4.7 percent of live births). Only one-fifth of these were identified from the birth certificate portion of the file. The hospital discharge diagnosis file was the largest contributor of unique diagnoses, providing over 65 percent of them. The distribution of the number of malformations per infant was 79 percent with only one malformation, 14 percent with two malformations, and 7 percent with three or more malformations.

The prevalence of smoking and malformations by selected population characteristics is shown in Table 1. Blacks reported having smoked during pregnancy more than Whites, the women aged 18–19 years reported smoking more often than the women in the older aged strata, the least educated mothers reported smoking the most, single mothers smoked more than married mothers, and the high multiparity mothers had a higher prevalence of smoking than lower parity mothers.

The adjusted odds ratios (OR) for all malformations for smokers of less than a pack a day vs nonsmokers was 0.99 (95% confidence interval (CI) = 0.93, 1.04). For smokers of one or more packs per day vs nonsmokers, the OR = 0.98 (95% CI = 0.92, 1.05). Because no dose response relation was observed in this data set between the level of smoking and all malformations, we chose to report models with only two levels of smoking (nonsmokers vs smokers).** For certain specific malformations, however, we did examine for a dose response relation across the three levels of smoking and report these results in the text.

**Results of other models available on request to authors.

TABLE 1—Distribution of Population Characteristics, Prevalence of Smoking and Prevalence of One or More Congenital Malformations among 288,067 Singleton Births to Missouri Residents, 1980–83

Population Characteristics	Percent of Population	Percent of Smokers in Each Group	Percent with Malformations in Each Group
Total Population	100.0	30.3	3.7
Race			
White	85.2	29.4	3.5
Black	14.8	35.5	3.7
Maternal Age (years)			
<18	5.7	36.0	3.8
18–19	10.0	41.9	3.6
20–24	34.6	35.5	3.5
25–29	31.0	24.4	3.5
30–34	14.5	21.8	3.5
≥35	4.2	25.7	4.2
Maternal Education (years)			
<12	32.2	52.2	3.6
12	22.7	30.0	3.5
>12	45.1	15.4	3.5
Marital Status			
Married	81.4	26.5	3.8
Single	18.6	47.0	3.9
Parity			
Primipara	36.9	27.4	4.0
Low multipara	43.1	29.0	3.7
High multipara	20.0	38.6	4.3

No congenital malformations occurred significantly more frequently in the infants of smokers than nonsmokers following adjustment for other maternal risk factors (Table 2). Only for gastrointestinal malformations did maternal smoking suggest a trend toward being a risk factor (OR = 1.11, 95% CI = 0.96, 1.29). Several congenital malformations, however, did appear to occur less frequently in infants of smokers. The odds ratio for malformations of the integument was less than 1.0 for smokers (OR = 0.82, 95% CI = 0.70, 0.97). The odds ratio for eye/ear/nose/throat malformations (OR = 0.84), ventricular septal defects (OR = 0.87), valvular defects (OR = 0.74) and cleft lip/cleft palate (OR = 0.84) suggested a lower prevalence in smokers although the 95% confidence intervals included 1.00.

We analyzed for several specific malformations within major organ systems that had been reported previously as being associated the maternal smoking during pregnancy. For example, the adjusted odds ratio for anencephaly for smokers compared with nonsmokers was 0.77 with a 95 percent CI of 0.44, 1.36. The odds ratio for atrial septal defects was 1.04 (95% CI = 0.70, 1.55).

The large number of odds ratios near 1.00 might suggest the data or methods used are not capable of detecting an effect of smoking. As an example of the validity of our model, we examined the odds ratio for chromosome malformations for various age groups. It is widely accepted that women over 35 years of age are at higher risk for having an infant with a chromosome abnormality. In fact, our data show the adjusted odds ratio for a chromosome malformation for women age 35 years and older was 7.04 (95% CI = 4.82, 10.28). There was no significant age-parity interaction for chromosome malformations ($p > 0.5$).

Discussion

The reported incidence of congenital anomalies varies from a low of 0.8 percent of live births[16] to a high of 7.1 percent.[17] This variation in the reported incidence of malformations makes it extremely difficult to determine the

TABLE 2—Adjusted Odds Ratios[a] for Congenital Malformations for Smokers Among Singleton Births to Missouri Residents by Congenital Malformation, 1980–83

Congenital Malformation	Number of Malformations	Adjusted Odds Ratio	95 Percent CI
All Malformations	10,223	0.98	0.94,1.03
Central Nervous System	489	1.04	0.85,1.26
Anencephaly	67	0.77	0.44,1.36
Eye/Ear/Nose/Throat	533	0.84	0.68,1.02
Heart	1,341	0.92	0.82,1.05
Ventricular septal	447	0.87	0.70,1.09
Atrial septal	123	1.04	0.70,1.55
Valvular	151	0.74	0.50,1.09
Other cardiovascular	756	1.01	0.86,1.19
Respiratory	299	1.01	0.83,1.24
Cleft Lip and/or Palate	451	0.84	0.68,1.05
Gastrointestinal	961	1.11	0.96,1.29
Genitourinary	1,622	0.97	0.86,1.08
Musculoskeletal	3,705	0.99	0.92,1.06
Clubfoot	980	1.02	0.88,1.17
Integument	830	0.82	0.70,0.97
Chromosome	330	0.99	0.76,1.28
Residual	435	0.95	0.77,1.18
Umbilical hernia	80	1.09	0.62,1.65

[a]Odds ratios adjusted for maternal age, race, martial status, educational attainment and parity

validity of reports of associations between congenital malformations and potential risk factors. Because the Missouri Birth Defect File uses several sources of information to obtain diagnoses of malformations, we feel this file comes close to accurately assessing the number of malformations in a birth cohort. In addition, ascertainment of malformations is more complete because two files, the Crippled Children Services file and the death certificate file, include diagnoses made up to two years after birth. The Missouri Birth Defects Registry's completeness is similar to the Centers for Disease Control Birth Defects Monitoring Program.[18] For many birth defects, the Missouri system appears to be more complete. However, the Missouri system is not as complete as the more comprehensive Metropolitan Atlanta Congenital Defects Program.[18]

In our analysis we did not find a relation between maternal smoking during pregnancy and the occurrence of congenital malformations. These findings merit attention because of the large sample size available for this analysis and the opportunity to adjust for variables that may confound the relation between smoking and malformations. For example, the positive associations with smoking reported by Shiono, *et al*.[19] for ventral hernias (OR = 10.1), omphaloceles (OR = 3.0), and other gut abnormalities (OR = 12.6) are based upon five, five, and six cases, respectively. Shiono's results may well be based on the chance distribution of a larger number of cases among smokers.

Other studies where positive associations have been reported between maternal smoking and congenital malformations may suffer from problems of failure to adjust for potential confounding variables.[20,21] Evans, *et al*,[21] reported an unadjusted association between maternal smoking and anencephaly (relative risk = 1.76 for smokers of more than 20 cigarettes per day). When the relation was examined by socioeconomic status the dose response relation between smoking and anencephaly was not apparent in the lowest socioeconomic strata and was inconsistent in the upper strata. Both socioeconomic status and race have been reported to be associated with congenital malformations as well as smoking.[17,22,23]

Several reports have suggested a lower rate of malformations in infants of smokers.[19,24,25] We did observe one group of malformations (malformations of the integument) for which maternal smoking during pregnancy did seem marginally protective (OR = 0.82, 95% CI = 0.70, 0.97). This observation was also reported by Shiono, *et al*.[19] Hook, *et al*,[25] reported a lower risk of Down Syndrome in infants of smokers. We observed no protective relation between smoking and chromosomal abnormalities (OR = 0.99).

Our analysis of the relation between maternal smoking during pregnancy and the occurrence of congenital malformations does not support a causal role for smoking. The validity of this observation may be questioned on the basis of an under-ascertainment of smoking using birth certificate data attenuating the relation between malformations and smoking. However, the prevalence of smoking in the Missouri data is similar to that reported in other populations.[3] In addition, we have no reason to believe there was differential reporting of smoking between women who had infants with malformations and those who did not have malformed infants (especially since many of the malformations were identified from sources other than the birth certificate). Secondly, the validity of our analyses might be questioned on the grounds of lumping specific malformations into groups of malformations within organ systems. However, the analysis of many more specific malformations with many fewer cases results in lower statistical power to find a relation and greater chances of finding spurious relations. We did analyze for several specific malformations that others had found to be related to maternal smoking and found no relation (e.g., anencephaly, ventricular septal defects, atrial septal defects, valvular defects, clubfoot, and umbilical hernias). It is interesting to note that the lumping of specific malformations did not mask the risk associated with age for certain malformations, e.g., Down Syndrome, within the chromosome malformation group. Within our model for chromosomal anomalies we did confirm the marked increased risk associated with age over 35 years for both multiparous women and primiparous women (OR = 7.04). This observation may lend support to the validity of analyzing the data across groups of malformations.

The lack of relation between smoking during pregnancy and congenital malformations in liveborn infants may be partially a result of a selection bias. Because smokers have twice as many abortions as nonsmokers[26] and 30 percent more fetal deaths,[3] fetuses with malformations among smoking mothers may be underrepresented in a study of live births. Despite the lack of a relation between maternal smoking during pregnancy and the occurrence of malformations in liveborn infants, the established relation between smoking and infant mortality and morbidity[3,4,27] argues against maternal smoking during pregnancy.

APPENDIX

International Classification of Diseases Ninth Revision (ICD-9) Codes Defining Malformation Categories

Cause of Death	ICD-9 Codes
Central Nervous System	7400-7429
Anencephaly	7400
Eye/Ear/Nose/Throat	7430-7449
Heart	7450-7469
Ventricular septal defects	7454
Atrial septal defects	7455
Valvular defects	7460-7467
Other Cardiovascular Defects	7470-7479
Respiratory	7480-7489
Cleft Lip and/or Palate	7490-7499
Gastrointestinal	7500-7519
Genitourinary	7520-7539
Musculoskeletal	7540-7569
Clubfoot	7545-7547
Integument	7570-7579
Chromosome	7580-7589
Residual*	7590-7599
Umbilical Hernia	5521, 5531

* Residual—remaining diagnoses in the ICD-9 Congenital Malformation Chapter (7400-7599).

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Conference on Health and Social Services for Elderly, Disabled Announces Call for Abstracts

The Department of Social Security of Regione Emilia Romagna is organizing an international scientific meeting, to be held in Bologna, Italy from April 17–21, 1990, to address the systems and science related to health and social services for the elderly and the disabled.

The objective of the conference is to focus on the problems of planning, organizing, and managing health and social services for this population, utilizing a multidisciplinary approach to solving the complex problems. Conference participants will be drawn from various groups, including systems analysts and scientists, professional health care providers, and policy-makers so that the final program will emphasize the interactions of the various interests and disciplines.

Interested program participants may submit an abstract (maximum one page, typed) representing an individual presentation, or proposals for an entire session, with subject matter and individual speakers identified. The deadline for submitting abstracts and session proposals is December 15, 1989. Papers accepted by the Scientific Committee will be published in the conference proceedings.

Send abstracts and proposals to: Professor Pier Paolo Puliafito, Università di Genova, Viale Causa, 13, 16145 Genova, Italia. Telefax: 39-10-310086.

Does Maternal Employment Affect Breast-Feeding?

NATALIE KURINIJ, PHD, PATRICIA H. SHIONO, PHD, SANDI F. EZRINE, MS, AND GEORGE G. RHOADS, MD, MPH

Abstract: A prospective survey of maternal employment and breast-feeding initiation and duration was conducted among 668 Black and 511 White women who delivered their first child in Washington, DC. Ninety-one percent of White women (n = 511) and 80 percent of Black women (n = 668) reported working during pregnancy. Black women who planned to return to work part time vs full time were more likely to breast-feed rather than formula-feed (adjusted odds ratio, 2.3; 95% confidence intervals (CI) = 1.4, 3.7). Using Cox regression, Black women who returned to work had a shorter duration of breast-feeding than those not returning to work (hazard ratio = 0.5 (CI = 0.3, 0.9)). Black and White women returning to professional occupations had a longer duration of breast-feeding compared to women returning to sales or technical positions (hazard ratio for Black women = 2.4 (CI = 1.4, 4.4); hazard ratio for White women = 1.6 (CI = 1.0, 2.5)). In addition, White women in professional occupations had a longer duration of breast-feeding than women in clerical positions (hazard ratio = 1.7 (CI = 1.1, 2.6)). Until employers in the United States develop a maternity policy which does not discourage breast-feeding, the recommended six months of breast-feeding will be difficult to achieve for most employed women. (*Am J Public Health* 1989; 79:1247–1250).

Introduction

Participation of women with infants in the labor force has risen dramatically over the last 10 years. In 1977, 32 percent of women with a child under one year of age worked outside the home[1] By 1982, this proportion increased to 43 percent, and today the majority (52 percent) of women with infants are in the labor force.[1] While employment could adversely affect the initiation and continuance of breast-feeding, there is little evidence to document this relationship.

In this analysis, we examine the relation between maternal employment and the initiation and duration of breast-feeding separately among Black and White women. We assessed whether maternal employment during pregnancy and the intention to return to work affect a woman's decision to breast-feed, whether the duration of breast-feeding differs between women who return to work and those who do not, and the employment variables (occupation, timing of return to work, or hours worked) which are associated with duration of breast-feeding.

Methods

Healthy Black and White primiparae over 17 years of age who delivered a normal, singleton infant weighing more than 2200 grams were consecutively selected from the delivery logs at each of three metropolitan Washington, DC, hospitals. The study design and procedures have been described.[2] A total of 1,409 women were asked to participate in the study and 84 percent (n = 1,179) agreed. Women were interviewed in the hospital between February 1984 and March 1985. Information was obtained regarding maternal sociodemographic characteristics (age, education, marital status, and occupation during pregnancy), the timing of anticipated return to work, as well as whether this would be full-time or part-time work. Breast-feeders were defined as women who breast-fed at all regardless of formula supplementation (n = 755).

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Women who breast-fed in the hospital (n = 755) were subsequently interviewed in their homes at one, three, and seven months postpartum. Any women who indicated she had stopped breast-feeding at the one-month or three-month home interview was subsequently interviewed by telephone with a shortened interview schedule. Women still breast-feeding at seven months postpartum were contacted by phone at one year to determine their infant-feeding history. Thus, we have information on duration of breast-feeding from birth to one year postpartum and information on maternal employment from birth to 7 months postpartum.

Of the 755 women choosing to breast-feed, 36 who started school were deleted because there were too few for meaningful comparisons with employment. Follow-up interviews among the 718 women (306 Black, 412 White) provided information on whether a woman was still breast-feeding, and if not, when she had stopped, when a woman returned to work, the number of hours she worked and her current occupation.

The US Census Bureau categories were used to define occupation codes.[3] Managers, writers, lawyers, and registered nurses were the most frequent occupations in the professional category. The clerical group was comprised mostly of secretaries and stenographers. The service professions were primarily represented by private household occupations, and occupations providing food, health, or personal services. Sales persons were largely involved in the sale of personal goods and services, and technicians were mostly in fields other than health or science.

Analyses

Logistic regression was used to estimate the adjusted odds ratio for breast-feeding while controlling for sociodemographic and employment variables. Coefficients were estimated by the maximum likelihood method.[4] The Cox proportional hazards regression model[4] was used to determine the employment variables associated with breast-feeding duration after controlling for socioeconomic variables. Life-table analysis was used to graphically present the effect of occupation on duration of breast-feeding.[4]

Results

Initiation of Breast-feeding and Anticipated Return to Work

Eighty percent of Black women (n = 668) and 91 percent of White women (n = 511) reported working during pregnancy (Table 1). Black women worked primarily in clerical

occupations, White women in professional occupations. Women who worked in professional or technical occupations during pregnancy were more likely to breast-feed than women in other occupations or those not employed. Breast-feeding rates did not differ between women who anticipated a return to work in the first year and those who did not, with or without adjustment for the effects of maternal education, age, and marital status in a multivariate model.

Among the 560 Black women and 371 White women who stated they intended to return to work, 71 percent of Black women and 82 percent of White women did so by the fourth postpartum month. Among White women, breast-feeding rates were significantly lower among women planning a return to work in the first two months postpartum than among those planning a later return (Table 1). This effect was present to a lesser extent among Black women. Hours of planned work was not associated with breast-feeding rates in either ethnic group. However, a greater proportion of Black women (74 percent) compared to White women (61 percent) anticipated a return to full-time work (more than 24 hours a week).

Logistic regression analysis was conducted in each ethnic group to determine which employment variables affected choice of infant feeding method in hospital. After adjusting for maternal age, education, marital status, occupation, and planned timing of return to work; the only employment variable significantly associated with breast-feeding among Black women was a planned return to part-time vs full-time work, OR 2.3 (CI = 1.4, 3.7). In a similar analysis among White women, no maternal employment characteristic was significantly related to breast-feeding in the hospital.*

TABLE 1—Incidence of Breast-Feeding in the Hospital by Maternal Anticipated Return to Employment†

Characteristic	Black (n = 668) No.	Black (n = 668) % Breast-feeding	White (n = 511) No.	White (n = 511) % Breast-feeding
Occupation during pregnancy				
Professional	68	82	237	92
Technical	21	76	26	96
Sales	73	41	39	77
Clerical	278	50	122	77
Service	95	37	37	78
Student	54	37	7	100
Homemaker	16	50	24	88
Unemployed	62	31	17	35
P-value		***		***
Planned to return to work within one year of delivery				
Yes	560	49	371	84
No	88	43	91	86
Don't know	20	55	49	84
P-value				
Month after delivery planned to return to work				
1 to 2 months	187	43	138	77
3 to 4 months	214	54	166	87
5 to 12 months	114	47	55	91
Don't know	44	51	12	92
P-value				*
Number of hours/week planned to work				
Full-time (>24 hrs)	413	47	207	83
Part-time	118	54	144	85
Don't know	26	65	20	90
P-value				
Maternal age (years)				
18–25	451	39	140	68
26–30	154	66	177	87
≥31	62	71	194	94
P-value		***		***

*p < .05, ***p < .001.
†The p values are derived from chi-square statistics comparing breast- vs formula-feeding.

Duration of Breast-feeding and Maternal Return to Work

Women returning to professional occupations had a longer mean duration of breast-feeding than women in other occupations (Table 2). Among White women only, the later the return to work the longer the mean duration of breast-feeding. Mean duration also was longer for women returning to part-time rather than full-time work.

Cox proportional hazard regression analyses in each ethnic group are presented in Table 3 and were conducted to determine if a woman's return to work and the timing of this return influenced duration of breast-feeding. After adjustment for the effects of demographic variables and formula use in the hospital, the hazard ratio (HR) of 0.5 for Black women who did not return to work compared to Black women who return between five and seven months postpartum indicates that the former were twice as likely to continue breast-feeding as the latter.

Cox proportional hazard regression analyses were also conducted to examine the effect of employment variables on duration of breast-feeding among women who returned to work (Table 4). Black women who worked in sales or technical occupations and White women who worked in clerical occupations discontinued breast-feeding sooner than women in professional occupations (Table 4). In addition, White women in sales or technical occupations had a shorter duration than professionals, HR = 1.6 (CI = 1.0, 2.5). There was a trend toward a shorter breast-feeding duration for women who returned to work full-time rather than part-time.

Life-table analysis by maternal occupation is presented in Figure 1. Duration of breast-feeding was longest among women in professional occupations; one-half of them were still breast-feeding at 38 weeks postpartum. In the other occupational groups, one-half were still breast-feeding at 14 to 22 weeks. The effect of maternal occupation on duration of breast-feeding was, for the most part, similar in both ethnic groups.

Discussion

A 1981 national survey found that college-educated women, who did not work outside the home, were more likely to continue breast-feeding than those who were employed.[5] Another survey of a selected sample of primarily White, married women[6] found that the timing of a woman's return to work and the number of hours she worked were important determinants of weaning while specific occupation had little effect. With the exception of these two reports, there is little empirical data regarding the effect of employment on breast-feeding.

We previously examined the effects of maternal age, education, ethnicity, hospital of delivery, family income, marital status, and attendance in childbirth classes on the incidence of breast-feeding in this sample of women.[2] Initiation of breast-feeding was most strongly associated with increasing maternal education.[2] The odds ratio for breast-feeding increased from 2.6 (CI = 1.9, 3.7) for women with 13 to 16 years of education to 5.2 (CI = 2.7, 10.2) for those with greater than 17 years of education when compared to women

*Data available on request to author.

TABLE 2—Mean Duration (in weeks) of Breast-Feeding by Sociodemographic and Employment Characteristics in Black and White Women*

Characteristics	Black			White		
	No.	Mean	(SD)	No.	Mean	(SD)
Occupation						
Sales	25	6.3	(6.8)	22	24.6	(20.0)
Technical	8	9.0	(6.6)	21	14.1	(12.2)
Clerical	109	11.6	(12.6)	52	17.5	(17.8)
Service	26	16.5	(14.5)	30	20.8	(20.0)
Professional	42	21.0	(17.0)	159	31.7	(18.1)
Actual return to work (months postpartum)						
1 to 2	88	12.2	(12.8)	99	22.0	(19.0)
3 to 4	103	14.6	(14.4)	144	26.9	(18.6)
5 to 7	41	13.0	(16.2)	56	32.2	(19.1)
Did not return by 7	74	16.9	(20.6)	113	30.4	(21.4)
Work hours						
Full-time	198	12.6	(13.1)	193	22.6	(18.2)
Part-time	32	19.2	(18.7)	103	32.6	(18.9)
Education (yrs)						
≤12	115	10.8	(13.3)	57	13.2	(14.6)
13–16	156	15.4	(16.7)	186	28.2	(20.4)
17+	35	20.8	(18.0)	169	31.3	(18.7)
Age (yrs)						
18–25	161	12.2	(14.4)	91	19.8	(19.0)
26–30	101	14.9	(16.5)	142	25.9	(19.7)
31+	44	20.3	(18.7)	179	32.5	(19.0)
Married						
No	147	12.9	(16.0)	24	12.5	(13.5)
Yes	159	15.6	(15.9)	388	28.3	(19.8)
Used formula in hospital						
Yes	163	11.9	(15.3)	112	22.5	(20.2)
No	143	17.0	(16.3)	300	29.2	(19.4)
Total	306	14.3	(16.0)	412	27.4	(19.8)

*Duration of brest-feeding for women continuing to breast-feed at 52 weeks postpartum was set to 60 weeks (n = 91).

TABLE 3—Proportional Hazard Ratios for Stopping Breast-Feeding among Black and White Women

Variables	Black (n = 306)		White (n = 412)	
	Ratio	95% CI	Ratio	95% CI
Actual return to work (months postpartum)				
5–7	1.0	—	1.0	—
≤ 2	1.0	0.7, 1.6	1.5	0.98, 2.2
3–4	1.1	0.7, 1.7	1.3	0.9, 1.9
Did not return by 7	0.5	0.3, 0.9	0.8	0.5, 1.2
Maternal education	0.9	0.86, 1.02	0.9	0.87, 0.97
Formula supplements in the hospital				
No	1.0		1.0	
Yes	1.5	1.1, 2.0	1.4	1.1, 1.8
Maternal age	1.0	0.9, 0.99	1.0	0.9, 0.99
Married				
No	1.0	—	1.0	—
Yes	1.1	0.8, 1.4	0.5	0.3, 0.9

TABLE 4—Proportional Hazard Ratios for Stopping Breast-Feeding among Black and White Women Who Returned to Work

Variables	Black (n = 208)*		White (n = 281)*	
	Ratio	95% CI	Ratio	95% CI
Occupation				
Professional	1.0	—	1.0	—
Clerical	1.2	0.8, 1.9	1.7	1.1, 2.6
Service	0.7	0.4, 1.4	1.1	0.7, 2.0
Sales/Technical	2.4	1.4, 4.4	1.6	1.0, 2.5
Return to work (months postpartum)				
5–7	1.0	—	1.0	—
≤ 2	1.2	0.7, 2.1	1.5	0.97, 2.3
3–4	1.3	0.7, 2.2	1.4	0.9, 2.1
Hours worked				
Part-Time	1.0	—	1.0	—
Full-Time	1.6	0.95, 2.7	1.4	1.0, 1.9
Maternal education	0.9	0.84, 1.02	0.9	0.86, 1.02
Formula supplements in the hospital				
No	1.0	—	1.0	—
Yes	1.8	1.3, 2.6	1.4	1.0, 1.9
Maternal age	0.9	0.89, 0.98	1.0	0.95, 1.02
Married				
No	1.0	—	1.0	—
Yes	1.1	0.8, 1.6	0.5	0.3, 0.8

*Sample size is decreased from 232 in Black and 299 in White women due to missing data.

with a high school education or less. Ethnicity was also associated with the initiation of breast-feeding (OR for White vs Black = 2.0 (CI = 1.4, 3.1)).[2] Duration of breast-feeding was substantially shorter among Black as compared to White women.[2] In both ethnic groups, the use of formula supplements in hospital was associated with a shorter duration of breast-feeding.[2]

In this paper we examined the effect of employment on initiation and duration of breast-feeding. Black women who anticipated a return to work within the first postpartum year and intended this return to be part-time were twice as likely to breast-feed in the hospital as those planning to work full time. Among White women, neither the planned timing of return to work, planned hours of work, or occupation were related to breast-feeding in the hospital. The majority of Black women (63 percent) planned a return to full time work, and the breast-feeding rate was lowest in this group. The socioeconomic need to return to work in the early postpartum months on a full-time basis may be strong among Black

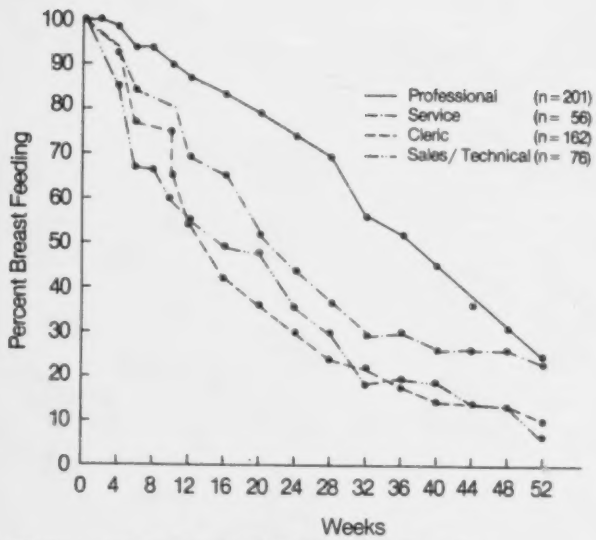

FIGURE 1—Duration of Breast-Feeding by Maternal Occupation.

women in this study who were largely single mothers. This need to return to full-time work may have had an adverse effect on the initiation of breast-feeding.

The effect of actual return to work was also different in the two ethnic groups. Black women who had not returned to work at seven months postpartum breast-fed longer than those returning to work earlier; there was, however, no difference between Black women returning to work early, e.g., at one or two months postpartum, and later, e.g., at six or seven months postpartum. Among White women, there was a positive trend between discontinuing breast-feeding and month of return to work; however, this trend was not statistically significant.

Occupation influenced duration of breast-feeding, and this association remained after adjustment for maternal sociodemographic characteristics. Women in professional occupations had a significantly longer duration of breast-feeding than women in sales or technical occupations. White women in professional occupations breast-fed longer than those in clerical occupations.

Because employers in the United States do not follow a general maternal leave policy, many working women by necessity resume employment during a time period that conflicts with the need for considerable mother-infant contact in breast-feeding. Expressing breast milk is possible but difficult in many work situations. The findings of this study emphasize that women employed as professionals, appear to have control over their work environment and can structure a more satisfactory relationship between the demands of employment and infant feeding. On the other hand, women employed in clerical and sales occupations appear to have little freedom to accommodate the demands of employment and the needs of their infants. Breast-feeding for six months postpartum is a recommended goal[7] which is difficult to meet for most employed women and particularly difficult for those employed in non-professional occupations. This finding emphasizes the need for the development and implementation of a maternal leave policy, particularly for women who choose to breast-feed.

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AACC Press, the publishing arm of the American Association for Clinical Chemistry, is actively seeking new listings for the 1990/91 edition of the *Directory of Rare Analyses* (DORA). The directory, currently in its third edition, is a vital reference tool for physicians and others seeking analyses not commonly performed in hospital or reference laboratories. The directory contains detailed listings of these uncommon tests, along with the address and contact person for the laboratory performing the test.

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Pregnancy Characteristics and Outcomes of Cambodian Refugees

PETER GANN, MD, MS, LUAN NGHIEM, BS, AND STANLEY WARNER, MD

Abstract: This study describes the perinatal characteristics of Cambodian refugees in Massachusetts. Data were abstracted from the records of 452 consecutive pregnancies among Cambodian women and 110 low-income Whites receiving obstetrical services at the same clinic and hospital in Lowell, Massachusetts. Pregnancies of Cambodian women were marked by a higher proportion of older mothers, grand multiparas, previous adverse birth outcomes, and short interpregnancy intervals. We identified maternal anemia (29.9 percent with hemoglobin <110 g/L) and inadequate utilization of prenatal care (32.3 percent with first visit in the 3rd trimester) as possible risk factors for the Cambodians. The prevalence of primary cesarean birth was only 6.3 percent in the Cambodians, compared to 15.6 percent in the comparison group, largely due to the infrequent occurrence of prolonged labor among multiparas. Despite the prominence of several risk factors for adverse birth outcomes in this population, major pregnancy complications were less common and the prevalence of low birthweight (6.4 percent) was close to the state average. Logistic regression analysis of risk factors for low birthweight identified young maternal age and short stature as the strongest factors operative in this community. Many of our findings are consistent with a strong cultural emphasis on managing the size of the baby to avoid a difficult labor and delivery. (*Am J Public Health* 1989; 79:1251–1257.)

Introduction

Since 1981, refugees from Cambodia have become a growing part of the flow of immigrants from Southeast Asia to the United States. These refugees are primarily rural people who are marked by the experience of having survived mass exterminations in Cambodia in the late 1970s and the harsh conditions of relocation camps in Thailand.[1] Not surprisingly, the emphasis in this refugee community on the replenishment of their population with healthy children is reflected in a high birth rate.[2] The maintenance of a cohesive culture with markedly different attitudes toward pregnancy and childbearing could provide insights into the complex relation between cultural beliefs, poverty, and pregnancy outcome.

This paper describes a survey of pregnancy characteristics and outcomes among Cambodian refugees in Lowell, a small city in northeastern Massachusetts. Lowell has a total population of 97,000 that now includes approximately 16,000 Cambodian refugees. An additional 7,000 were expected to arrive during 1988, attracted by entry-level jobs in the electronics and service industries and an established base of family networks and cultural organizations. Although the rate of employment among these refugees is relatively high, low salaries and educational levels as well as crowded living quarters are serious problems. According to data compiled by the US Centers for Disease Control, the prevalence of low birthweight in two of the largest relocation camps in Thailand was at least 20 percent.[3] Three published studies, which included relatively small numbers of Cambodian pregnancies, suggested that Cambodians had less favorable birth outcomes than other Southeast Asian groups such as the Hmong and Vietnamese.[2,4,5] A study of mental health problems among all Indochinese refugees concluded that Cambodian women, particularly the many who had lost spouses, had the highest risk for serious psychiatric impairment.[6] We initially assumed that current living conditions—in addition to the trauma and deprivation of the past, history of poor pregnancy outcomes, and the difficulty of relocating to a strange culture—would put the Cambodians at risk during pregnancy and childbirth in the US.

The purpose of this study was to define the prevalence of adverse birth outcomes and to describe the major risk factors in this large population. In addition, we contrasted the pregnancy characteristics of this relatively unique racial/ethnic group with those of White, native-English speaking women of low income in Lowell and in Massachusetts as a whole.

Methods

Lowell General Hospital (LGH) is the only hospital in Lowell that provides Cambodian interpreters for prenatal patients and is currently the only hospital in Lowell providing obstetrical care for Medicaid patients. Therefore, nearly all Cambodian women in Lowell who receive prenatal care will at one time or another during pregnancy become registered at LGH. Birth lists and a log of patients registered in the prenatal clinic were reviewed to select all Cambodian women receiving prenatal care at LGH between January 1, 1981 and July 30, 1987. Women transferred for care or delivery at another hospital, women who left the area during pregnancy, home births and stillbirths were recorded in the prenatal clinic log. Attempts were made to trace each of these women and obtain the obstetrical records. Only two women, who moved out of the area before delivery, could not be traced. A total of 449 pregnancies (444 singleton pregnancies) among 310 women were found. The number of pregnancies increased from 12 in 1981 in 119 in the first seven months of 1987.

A trained investigator reviewed the medical chart for each Cambodian pregnancy using a coding manual to abstract detailed information on 53 variables concerning the patient's demographic and medical features; none of these variables were missing for more than 3 percent of the eligible subjects. Identical information was obtained from the records of 110 comparison pregnancies that were registered at the same clinic immediately after every fourth Cambodian pregnancy. These patients, who represent the majority at the clinic, were Caucasian women (often of Greek, Irish, Italian, or French-Canadian ancestry) whose primary language is English. As in the Cambodian group, 90 percent of these patients were

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registered for Medicaid. For some variables, comparison data were also available for White or non-Asian women participating in the statewide Maternal and Infant Care (MIC) Program providing services for low-income women.[7] This group is 40 percent White, 27 percent Black, and 33 percent Hispanic. Comparison groups were selected to delineate contrasts due to cultural or genetic factors rather than socioeconomic ones.

We defined low birthweight as a birthweight less than 2500 grams and premature delivery as gestational age less than 37 weeks. Adequacy of prenatal care was determined using an established index developed by the Institute of Medicine.[8] Details of this index and other key definitions—such as gestational age, rate of maternal weight gain, labor times and inter-pregnancy interval—are provided in the Appendix. Confidence intervals (95%) were calculated using normal approximation or exact methods as appropriate. Crude and stratified odds ratios were calculated to estimate the magnitude of association between risk factors and outcomes. A microcomputer program specifically designed for logistic regression was used for multivariate analyses.[9]

Results

The prevalence of selected major maternal risk factors existing *prior* to pregnancy is shown in Table 1. There are fewer teenage mothers and primiparas, and more older mothers and grand multiparas than in the comparison group; nearly one-fourth of the Cambodian mothers were under 60 inches in height; a greater proportion had experienced previous adverse birth outcomes; and over one-fourth became pregnant six months or less after their previous delivery.

Table 2 shows the risk factors observed during pregnancy. Although the mean weekly weight gain was slightly less among the Cambodians, there was very little difference in the proportion of women experiencing sharply reduced rates of weight gain. On the other hand, the Cambodians were far less likely to experience high rates of weight gain (more than 1.5 lbs/week). Anemia, as determined by hemoglobin concentration, was more common among the Cambodian women. Cigarette smoking and regular alcohol consumption were far more common in the clinic comparison population. There was a small difference between the groups in prenatal care, but both Lowell populations were considerably worse in this regard than the statewide MIC population, in which only 14.3 percent received "inadequate" prenatal care.

Labor and delivery results are shown in Table 3. The major finding is the much lower prevalence of primary cesarean birth among the Cambodians, almost entirely attributable to a difference between the multiparous women. Prolonged first stage of labor and a clinical diagnosis of dystocia, the leading reason for primary cesarean section, were more common among multiparas in the comparison group. We detected a small increase in the occurrence of very short second stages of labor among the Cambodians, but the prevalence of potential precipitous labor was higher than expected in both groups.

The prevalence of adverse outcomes among groups in the study is shown in Table 4. The prevalence of low birthweight was lower in the Cambodians than the clinic comparison or statewide MIC women. The difference between Cambodians and the clinic group held for both primiparas and multiparas. The data in Table 4 and Figure 1 indicate that the frequency distribution for birthweight was narrower and slightly shifted to the left for the Cambodians compared to the clinic group. There were fewer stillbirths in the Cambodian population and premature birth was approximately equally common among all three groups. Comparison of gestational age distributions in Figure 2 reveals that the Cambodian women have a relative deficit of babies born after 40 weeks of gestation and an excess of babies born at marginally premature or early gestational ages from 35 to 39 weeks.

Table 5 presents the crude odds ratios for low birthweight and prematurity in the Cambodian population alone, considering several variables of interest. Although the effect estimates are not highly precise, there are some clear differences between the effect of certain variables on low birthweight versus premature birth. In particular, height, age less than 20, and nulliparity appeared to be more predictive of low birthweight than prematurity. Age less than 20 was in fact inversely associated with premature delivery. Slow maternal weight gain (less than 0.5 lbs/week) showed a relatively strong association with delivery of an underweight infant, a relationship that proved to be unconfounded by maternal stature.

Table 6 presents the results of a logistic regression analysis of the risk factors for low birthweight among the Cambodians. More saturated models gave similar point

TABLE 1—Maternal Risk Factors Prior to Pregnancy in Cambodian Refugee and Clinic Comparison Groups

Maternal Risk Factors	Cambodian Refugees n = 444	Clinic Comparison n = 110
Maternal Age		
>35 years	9.8% (7.0, 12.6)	1.8% (0, 4.3)
<20 years	13.2% (10.0, 16.4)	36.9% (27.9, 45.9)
Parity		
= 0	24.7% (20.2, 28.8)	50.4% (41.1, 59.7)
>4	11.6% (8.6, 14.6)	1.8% (0, 4.3)
Height		
mean	60.8" (60.5, 61.1)	68.3" (67.6, 69.0)
<60"	24.0% (20.0, 28.0)	0%
Previous LBW or premature birth*		
≥1	26.2% (20.0, 32.4)	10.0% (4.4, 15.6)
≥2	6.2% (2.8, 9.6)	2.7% (0, 5.7)
Previous infant death*	8.2% (4.4, 12.0)	2.7% (0, 5.7)
Interpregnancy interval ≤6 mos	25.4% (20.6, 30.2)	4.0% (0, 9.4)

95% confidence intervals in parentheses.
*Denominator includes first pregnancy in study only.

TABLE 2—Risk Factors During Pregnancy in Cambodian Refugee and Clinic Comparison Groups

Risk Factors	Cambodian Refugees n = 444	Clinic Comparison n = 110
Mean weekly weight gain	0.92 lbs. (0.87, 0.96)	1.06 lbs. (0.97, 1.15)
<0.5 lb/wk	12.2% (9.0, 15.4)	10.0% (4.1, 15.8)
>1.5 lb/wk	6.9% (4.4, 9.4)	18.0% (10.5, 25.5)
Anemia		
Hb <90 g/L	3.2%	0%
Hb 90–100 g/L	6.3%	1.8%
Hb 100–110 g/L	20.4%	12.0%
Total	29.9 (25.6, 34.2)	13.8 (7.3, 20.3)
Current smoker	3.3% (1.3, 5.3)	54.5% (45.2, 63.8)
Regular use of alcohol	1.3% (0, 2.6)	13.6% (7.2, 20.0)
First prenatal visit		
Week 0–13	10.9% (7.7, 14.1)	13.8% (7.3, 20.3)
Week 28+	32.3% (27.5, 37.1)	25.7% (17.5, 33.9)
Inadequate prenatal care*	34.9% (30.4, 39.4)	28.0% (19.5, 36.5)
Pregnancy complications†	19.8% (16.1, 23.5)	29.1% (20.6, 37.6)

95% CI in parentheses.
*See Appendix for definition.
†Includes such conditions as toxemia, gestational diabetes, genitourinary infection, and placental abnormalities.

TABLE 3—Labor and Delivery Characteristics of Cambodian Refugee and Clinic Comparison Groups

Characteristics	Cambodian Refugees n = 444	Clinic Comparison n = 110
1st stage of labor >12 hrs	9.2% (6.3, 12.1)	21.9% (13.0, 30.8)
Primiparas*	14.8% (7.9, 21.7)	15.1% (5.5, 24.7)
Parity 1–4*	7.3% (4.3, 10.3)	29.4% (14.1 44.7)
Dystocia	3.9% (2.1, 5.7)	7.3% (2.4, 12.2)
2nd stage less than 6 minutes		
Primiparas	6.0% (1.4, 10.6)	2.3% (0, 6.7)
Parity 1–4	43.0% (37.4, 48.6)	37.8% (22.2, 53.4)
Primary cesarean	6.3% (4.0, 8.6)	15.6% (8.8, 22.4)
Primiparas	12.1%	12.5%
Parity 1–4	4.3%	18.5
Repeat cesarean	1.6%	3.7%

*Primiparas: n = 108 (Cambodian), n = 56 (Clinic Comparison)
Multiparas: n = 282 (Cambodian), n = 54 (Clinic Comparison)
95% CI in parentheses.

TABLE 4—Prevalence of Adverse Birth Outcomes in Cambodian Refugee and Clinic Control Groups

Birth Outcomes	Cambodian Refugees	Clinic Controls	MIC, 1980–84
Low birthweight			All non-Asians: 9.1% (8.5, 9.7) White:
<2500 gms	6.4% (4.1, 8.6)	10.3% (4.6, 16.0)	7.3% (6.5, 8.1)
Primiparas	9.4%	12.7%	—
Multiparas	4.6%	7.7%	—
Mean birthweight	3146 gms	3326 gms	All non-Asians: 3215 gms White: 3291 gms
Birthweight 2500–3000 gms	28.3% (24.1 32.5)	15.9% (9.0, 22.8)	All non-Asians: 18.8% (18.0, 19.6)
Stillbirths	2 (0.4%)	3 (2.7%)	All non-Asians: 30 (0.7%)
Premature birth	9.1% (6.4, 11.8)	6.4% (1.8, 11.0)	All non-Asians: 8.2% (7.6, 8.7)

estimates for the effect of individual variables. Age less than 20 and maternal height less than 60 inches were the only substantial risk factors in the multivariate analysis. Adjusted odds ratios were generally similar to the crude odds ratios. Slow maternal weight gain was a strong independent predictor of low birthweight but was excluded from the multivariate

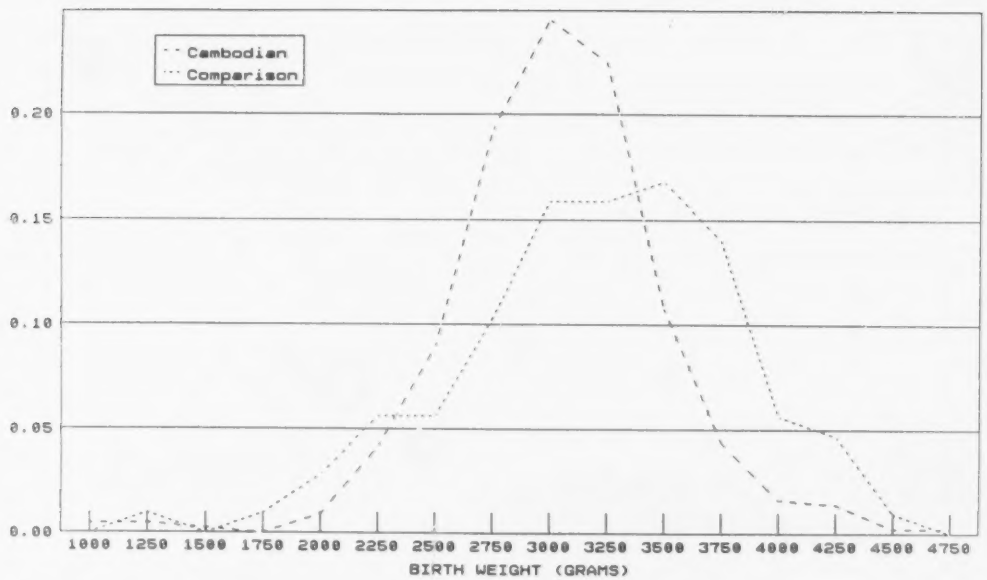

FIGURE 1—Relative Frequency of Birth Weights among Cambodian Refugees and Clinic Comparison Group

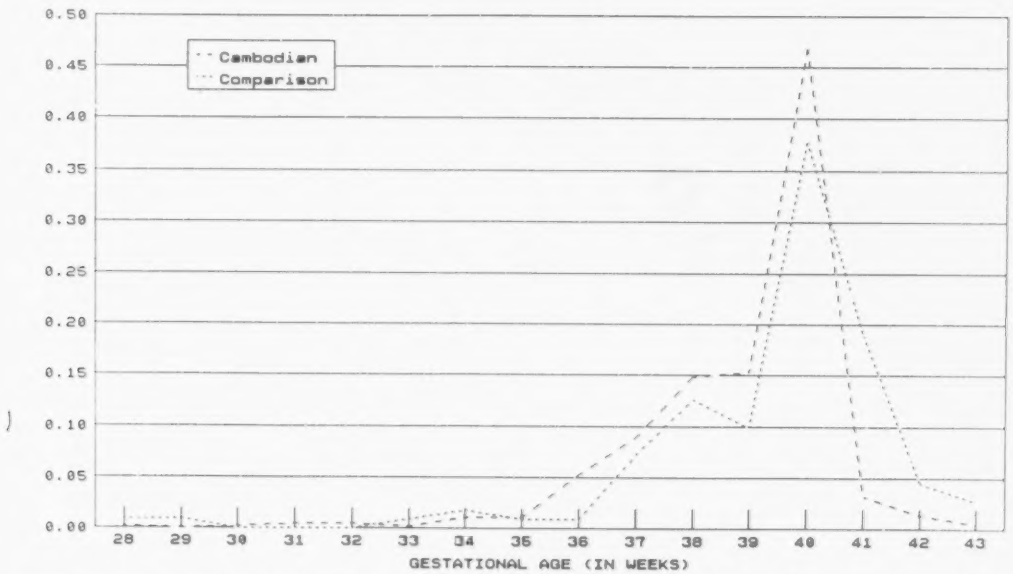

FIGURE 2—Relative Frequency of Gestational Ages among Cambodian Refugees and Clinic Comparison Group

models since it could be considered in part a manifestation of rather than a risk factor for a pregnancy impaired by intrauterine growth retardation.

Discussion

Contrary to our expectations prior to this study, the Cambodian refugees had more favorable birth outcomes than White low-income women in Lowell and low-income women in Massachusetts in general. The prevalence of low birthweight (LBW) infants, stillbirths, major pregnancy complications, and primary cesarean delivery were all lower in the Cambodian population. The prevalence of LBW was in fact similar to that of all non-MIC women in Massachusetts, which averaged 5.8 percent between 1980 and 1984. The prevalence of LBW among the refugees changed little during the interval from 1981 to 1987; it was 6.5 percent during 1981–84 and 6.2 percent during 1985–87. Some of the differences in the birthweight distribution between the refugee and other populations is attributable to more births taking place after 37–40 weeks gestation among the Cambodians. Jarvis and co-workers obtained a strikingly similar result when comparing birthweight and gestational age patterns of Indian and Pakistani immigrants to Whites in London.[10] The asso-

TABLE 5—Crude Odds Ratios for Low Birthweight and Premature Birth

Maternal Risk Factors	OR (95% CI) LBW	OR (95% CI) Premature
Height <60 inches	2.67 (1.20, 5.94)	1.12 (0.51, 2.49)
Slow maternal weight gain (<0.5 lb/wk)	3.08 (1.10, 8.70)	0.73 (0.21, 2.47)
Inadequate prenatal care	1.00 (0.71, 1.41)	0.89 (0.44, 1.83)
Interpregnancy interval <6 mos*	2.33 (0.77, 7.01)	1.09 (0.42, 2.80)
Medicaid payment	1.14 (0.42, 3.12)	1.51 (0.61, 3.71)
Previous LBW or premature birth*	1.25 (0.43, 3.64)	1.85 (0.84, 4.05)
Age <20 years	2.20 (0.85, 5.72)	0.17 (0.03, 0.90)
Age >35 years	1.48 (0.42, 5.24)	0.67 (0.20, 2.26)
Hemoglobin <100 g/L	2.38 (0.79, 7.18)	1.36 (0.45, 4.07)
Nulliparity	2.09 (0.89, 4.90)	0.83 (0.37, 1.88)
No previous US birth*	0.87 (0.36, 2.11)	1.78 (0.80, 3.98)

*Analysis includes multiparas only.

TABLE 6—Odds Ratios for Low Birthweight among Cambodian Refugees Adjusted by Logistic Regression

Risk Factor*	Odds Ratio (95% CI)
Age <20 years	3.43 (1.22, 9.60)
Height <60 inches	2.70 (1.04, 7.02)
Hemoglobin <100 g/L	2.41 (0.64, 9.13)
Inadequate prenatal care	1.19 (0.47, 3.05)

*Odds ratios for the major predictor variables were not significantly altered by inclusion of other independent variables in the model (see text).

ciation between low pre-pregnancy weight and slightly shorter gestation provides one possible explanation for such similar results with diverse Asian groups.[11]

Examination of the risk profile for the Cambodian women reveals several sharp contrasts to the other groups. Risk factors more common among the Cambodians (older mother, grand multiparity, previous small infants or infant deaths, short interpregnancy interval, inadequate prenatal care, and anemia) appear to have been offset at least in part by a low prevalence of teenage primiparous mothers and low rates of smoking and alcohol use. Muecke found similar contrasts in a comparison of Southeast Asian to non-Asian pregnancies in Seattle.[12] It is encouraging to note that the conditions leading to previous poor outcomes in this population, such as malnutrition or infection in the relocation camps, have not had an irreversible effect on pregnancy outcome in the US.

Short maternal stature was one of two major predictors of low birthweight in this population. Hardy and Mellits, using data from the Collaborative Perinatal Project, have shown that prevalence of LBW defined as 2500 grams or less begins to rise as maternal stature falls below about 63 inches.[13] It is reasonable to assume that many of the LBW infants born to these small women are physiologically sound, and probably reflect genetic tendencies toward small body size. Therefore, clinically significant LBW among these refugees might be overestimated in the present study. Previous investigators have pointed out the existence of ethnic variation in "normal" birthweights and the need for racially appropriate definitions of low birthweight based upon follow-up of morbidity and mortality during the first year of life.[14–17]

These data indicate that about one-third of the Cambodian refugee women enter prenatal care during the third trimester and about 35 percent receive inadequate prenatal care as determined by the timing and quantity of visits. This problem is also seen in the White patients from the same clinic, in comparison to low-income women statewide. Despite use of the hospital by a large number of Cambodians and the availability of interpreters on a limited basis, the Cambodians tend to view the hospital with some fear and/or mistrust during pregnancy. We found no evidence for any alternative, non-Western systems of prenatal service in the community, although it does appear that some Cambodian women have their pregnancy diagnosed relatively early during a single visit to one of the Vietnamese physicians in the community. We were able to analyze the use of prenatal care for succeeding pregnancies in Lowell and found that there was a slight tendency toward later prenatal registration with each succeeding pregnancy. Although we would not de-emphasize prenatal care for this group, we found, as others have in different populations, that the quantitative adequacy of prenatal care by itself was not an important factor in pregnancy outcome.[18]

Anemia is another problem that, although more frequent among the Cambodians, did not appear to have a substantial effect on birth outcome. Future studies will provide a more detailed analysis of the hematologic conditions in this population. Attention must be focused on possible dietary iron deficiency as well as chronic intestinal blood loss from parasitic infestation and hereditary hemoglobin disorders, which have been reported among Cambodian refugees by other investigators. Hemoglobin disorders that are likely to be found in this population include Hemoglobin E, and both alpha and beta varieties of thalassemia.[19] In a survey that included newly arrived Cambodian refugees, Catanzaro and Moser reported that 67 percent of the Cambodians had positive stools for intestinal parasites and that the presence of hookworm was associated with anemia.[20]

Previous investigators have postulated that Southeast Asian women have lower hematocrit and hemoglobin values during pregnancy due to physiologic adjustments having no clinical significance.[21] This could arise from greater expansion of plasma volume, for instance. This explanation appears to be consistent with our data showing no relation of anemia to birthweight until hemoglobin drops below 100 grams/L. Despite some relevant data from animal studies, the long-term effect of anemia during pregnancy on the child, particularly in regard to cerebral development, is not yet well understood.[22] The refugee's prenatal diet could be high enough in calories to avoid LBW, yet borderline in specific nutrients such as total iron, protein, and calcium.

Short interpregnancy interval was extremely common among the Cambodians. The number of LBW infants was too small to permit a reasonably precise, unconfounded estimate of the effect of interval on birthweight. However, we found no relation between short intervals and anemia or slow maternal weight gain in the second pregnancy. As pointed out by Klebanoff, interpregnancy interval could be more important in populations where maternal nutrition may not be adequate.[23]

The finding of a strong negative association between age under 20 and premature delivery is not typical in the general population and suggests that the LBW found among the Cambodian teenagers was specifically due to intrauterine growth retardation. The reasons for the apparent protection of teen mothers from premature delivery are not clear. Since teenage pregnancy may become more common among the Cambodians, the current low prevalence of LBW could begin to increase. The life experience and cultural practices of the

refugees under age 20, who came of age either in relocation camps or in the US, might differ considerably from that of the older women. So far, we have found no evidence for a trend toward an increase in smoking or alcohol use among the Cambodian teens. In fact, the few smoking women in our study tended to be older than the non-smokers.

Although the Cambodian women are small and tend to have small pelvic sizes, they had a lower frequency of labor complications and prolonged labor, resulting in a very low rate for primary cesarean delivery. The tendency for the Cambodians to have smaller babies, after shorter gestation, would explain at least part of this difference. We and others working with Southeast Asian refugees found a strong belief in the importance of limiting the size of the infant in order to avoid a difficult labor, but we found no evidence for any specific intervention by the mothers to achieve this other than some restriction of food intake during pregnancy. The fear of cesarean section seems quite reasonable given the risks of surgical delivery in rural Cambodian and the possible stigmatization of the cesarean newborn for religious or cultural reasons.

Among the strengths of the present study are the large size of the Cambodian population studied and the inclusion of virtually all pregnancies since 1981 in this population. In addition, the provision of nearly all obstetrical services by a single hospital resulted in a fairly uniform level of health care service. The study was limited by the surprisingly small number of low birthweight (LBW) outcomes and the absence of an LBW definition specific for this racial/ethnic group.

In summary, we found that the Cambodian refugees possess an unusual risk profile for adverse birth outcomes. Birth outcomes themselves were good in comparison to other groups of low-income women, leading us to conclude that previous poor outcomes in refugee camps were due to reversible conditions. Studies presently underway will determine whether these relatively good birth outcomes among Cambodian refugees in Massachusetts are complemented by lower rates of infant mortality. To an unknown extent, the outcomes observed in this group of women who survived the periods of extermination and relocation might also have been favorably influenced by the selection out of the reproducing population of women with chronic illnesses. Indeed, we found that relatively few of these women began pregnancy with major chronic diseases. Although strong additional interventions to improve basic outcomes may not be urgently needed, the problems of teen pregnancy, resistance to prenatal care, and the possible long-range effects of maternal anemia deserve particular attention in this community.

APPENDIX
Definitions

Gestational age: the final, best estimate of pregnancy duration based upon the last menstrual period, maternal examination, ultrasound, and clinical examination of the newborn.

First stage of labor: the interval from the onset of regular, painful contractions to the full dilatation of the cervix.

Second stage of labor: the interval from full dilatation to delivery of the infant.

Interpregnancy interval: the interval from the previous birth to the last menstrual period prior to the current pregnancy.

Rate of maternal weight gain: Since many of the women did not know their pre-pregnancy weight, it was often impossible to determine the total weight gain during pregnancy. Instead, the rate of maternal weight gain during pregnancy was estimated by dividing the weight obtained on the two most widely spaced clinic visits by the number of intervening weeks. This measure of average weekly weight gain is reasonable since the largest portion of the sigmoidal curve for weight gain during pregnancy is linear and observations could be discarded if the two weight measurements were too late or too close in time. We chose to discard values obtained for intervals less than four weeks or within six weeks of the due date.

	Gestation (weeks)	Number of Visits
Adequate[a]	22–25	4 or more
	26–29	5 or more
	30–31	6 or more
	32–33	7 or more
	34–35	8 or more
	36 or more	9 or more
Inadequate[b]	22–29	1 or less
	30–31	2 or less
	32–33	3 or less
	34 or more	4 or less
Intermediate	All combinations not specified above.	

[a]First prenatal visit must also be in the first 13 weeks of pregnancy.
[b]First prenatal visit after 28th week is automatically inadequate.

Adequacy of prenatal care index: This index (first published in reference 8) is as follows:

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New Curriculum Available to Upgrade Drug Abuse Counseling Skills

The National Institute on Drug Abuse has developed a new instructional training resource for employee assistance program (EAP) professionals and others who need to update or increase their knowledge and skills for responding to drug-using employees in the workplace.

Single copies of the document, entitled *Drug Abuse Curriculum for Employee Assistance Program Professionals*, are available free on request. It was developed in response to increasing concerns about the effects of drug abuse in the workplace and the need for EAP staff to be prepared to help employees with problems resulting from the abuse of a wide range of drugs. Several experts in the field—in consultation with NIDA and a national panel representing academic institutions, unions, businesses, EAP practitioners, professional associations, and the public sector—authored the modules of the *Curriculum*. The six modules include:

- Organizational Implications of Drug Abuse Programming: Making the Organization Work for You (by Dr. Sheila Akabas and Meredith Hanson);
- Supervisor Training (by Dr. Bradley Googins);
- Case Management (by Dr. Muriel Gray);
- Assessment (by Dr. Dale Masi);
- Overview of Drug Abuse and Employee Assistance Programs at the Workplace (by Dr. Judith Vicary); and
- Pharmacology of Abused Drugs (by Dr. George Woody).

The *Curriculum* covers the topics of overall manifestations of drug abuse in the workplace; pharmacology; the physical, psychological, and behavioral effects of drugs—both licit and illicit; the relationship of drug abuse to job performance and other workplace issues; the role of the EAP in drug abuse assessment and case management; supervisor training; and organizational responses.

The adaptability of the *Curriculum* allows for its use in both continuing education (CE) and in-service training programs offered through academic institutions and has been designed so that continuing education credits may be offered to participants. Pennsylvania State University has adopted the *Curriculum* for use in their CE Statewide Chemical Dependency Counseling Treatment Program which serves over 300 students annually. Other university-based CE programs are currently examining the *Curriculum* for potential adoption as a course text.

Single copies of *Drug Abuse Curriculum for Employee Assistance Program Professionals* are available free upon request from the National Clearinghouse for Alcohol and Drug Information, P.O. Box 2345, Rockville, MD 20852. Tel: 301/468-2600.

Comments or questions regarding the *Curriculum* should be directed to Jim Lipari, Chief, Employee Assistance Program Section, Workplace Policy Research Branch, Division of Applied Research, NIDA, Rockville, MD 20857. Tel: 301/443-0802.

Catholicism and Fertility in Puerto Rico

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Abstract: This paper examines the association between fertility and religious affiliation among Puerto Ricans. Bongaarts' proximate determinants framework is used to compare religious categories on behaviors that directly affect fertility. We found no difference in recent fertility between Catholics and non-Catholics in Puerto Rico (total fertility rate = 2.5). We also found Catholics and non-Catholics to be similar on age at first marriage (23.3 and 23.4 years, respectively), contraceptive use (71 percent and 69 percent of married women currently contracepting), and breast-feeding practices (mean duration of breast-feeding of 4.4 and 4.3 months). Differences were observed between less committed and more committed Catholics in total fertility rate, age at first marriage, and contraceptive use, but these differences were not large and were in the opposite direction from the expected. Our findings support the theory of convergence of Catholic and non-Catholic fertility for the US. Our discrepancy with findings from the National Survey of Family Growth is evidence of the cultural diversity in the US Hispanic population and indicates the importance of disaggregating by ethnicity or national origin when studying this population. (*Am J Public Health* 1989; 79:1258–1262.)

Introduction

There has been extensive study of the differences in Catholic and non-Catholic fertility in the United States over the past 30 years.[1–7] Early studies in the US found Catholics had higher fertility and lower contraceptive use than non-Catholics.[1,2] Later research presents two contrasting views. Some studies suggest Catholic and non-Catholic fertility behavior has converged to the point that the differential is negligible.[4] Other studies suggest Catholics continue to have distinctive fertility behavior relative to non-Catholics.[5–8] Most recently, Mosher, *et al*, have attributed the distinctiveness to the high fertility of Hispanic Catholics, such that when Hispanics were excluded from their analysis, Catholic and non-Catholic fertility in the United States converged. Among Hispanics, however, they observed an effect of religion, with Catholics having higher fertility than non-Catholics.[7] The analysis by Mosher, *et al*, however, is limited by the small sample of Hispanics in the National Survey of Family Growth and consequently an inability to disaggregate the Hispanic population into specific ethnic groups. Given the increasing numerical importance of the Hispanic population in the United States, if we are to draw definitive conclusions regarding Catholic and non-Catholic fertility in this country, we need studies that look at the Hispanic population controling for national origin.

Puerto Ricans are an important component of the US Hispanic population. Since Puerto Rico has commonwealth status with the United States, Puerto Ricans are United States citizens and have unrestricted travel to the mainland. They comprise the second largest Hispanic ethnic group residing in the United States, and their mobility is characterized by circular migration between Puerto Rico and the United States mainland. Because of the fluidity of movement between the island and the mainland, one can expect certain similarities between Puerto Ricans living in these two geographically distinct areas. Hence an analysis of the relation between Catholicism and fertility in Puerto Rico should have a direct bearing on this relation for Puerto Ricans living in the United States.

Our study will focus on both direct and indirect determinants of fertility in Puerto Rico, a framework of fertility analysis first introduced by Davis and Blake[9] and more recently modeled by Bongaarts.[10–12] Davis and Blake suggested that social, economic, and cultural factors (i.e., indirect determinants) operate through intermediate fertility variables (i.e., direct determinants) to affect fertility. Bongaarts, in modeling Davis and Blake's scheme, reduced the original set of 11 variables to four proximate direct determinants of fertility (nuptiality, contraceptive use, lactation, and abortion). Variations in the importance of the proximate determinants can either: account for group fertility differences, or explain how groups with similar fertility levels reached equality. In this paper, we use indicators for three proximate determinants (or intermediate variables) to examine the impact of religion on fertility in Puerto Rico. The central issue addressed is whether fertility differs by religious affiliation and/or religious participation in Puerto Rico. A further issue asks whether analysis of the intermediate fertility variables helps clarify the impact of religion on fertility in Puerto Rico. Two primary comparisons are made: 1) Catholic versus non-Catholic; 2) for Catholics the more committed versus the less committed.

Methods

The Sample

In 1982, the Puerto Rico School of Public Health and the Puerto Rico Department of Health, with technical assistance from the Division of Reproductive Health, Centers for Disease Control, conducted an island-wide assessment of fertility and family planning in Puerto Rico.

A two-stage, disproportionate stratified cluster sample design was used to select 4,500 households throughout the island of Puerto Rico. All women of ages 15–49 in each survey household were interviewed. A total of 3,175 women completed personal interviews, resulting in a completion rate of 91 percent. The data have been weighted to take into account: nonresponse at both the household and individual levels, and poststratification adjustments for age by residence, based on the age by residence distributions from the 1980 census of Puerto Rico and population estimates for Puerto Rico in 1982.

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The analysis presented in this paper uses weighted data. However, only unweighted Ns are shown in the tables.

Variables

A two-question sequence was used to define religion and religious commitment. (The complete version of these questions in Spanish and English are presented in the Appendix.)

The data set contains only four categories for religious affiliation—Catholic, Protestant, Other, and None. In this study, non-Catholics are treated as one group because of the small proportions in each of the three non-Catholic categories (Protestant = 16 percent, Other = 9 percent, and None = 7 percent).

As suggested by Mosher *et al*,[7] religious participation is measured by church attendance. Catholics who attended services once a month or more were categorized as frequent attenders or more committed Catholics and those who attended services less than once a month were defined as infrequent attenders or less committed Catholics. As a check on this indicator, analyses were also done using communion participation as a measure of religious commitment. Results using the latter variable paralleled those for church attendance but are not shown.

Based on the above considerations, four groups were classified: Catholics (2,144) and non-Catholics (1,031); and for Catholics, the more committed (1,204) and the less committed (940).

Two measures of fertility are used: 1) mean number of children ever born; and 2) total fertility rate, defined as the average total number of births a woman would have if she were to pass through her reproductive lifetime experiencing the age-specific fertility rates of a given time period. The total fertility rate for this analysis was computed using the cohort-period technique.[13]

Of the four proximate determinants of fertility described by Bongaarts (marriage, contraception, lactation, and abortion), this analysis addresses the first three. Abortion data are not included in this study because reliable estimates of abortion prevalence are not available. The relation between the proximate determinants and fertility are explained in detail in Bongaarts[10,11] and Bongaarts and Potter.[12]

Results

In the aggregate, the mean number of children ever born is the same for Catholics, regardless of commitment level, and non-Catholics (Table 1). When we look only at women aged 40–49 years, a small difference exists between Catholics and non-Catholics, suggesting that there may have been a mild effect of religion on childbearing at an earlier point in time regardless of commitment.

TABLE 1—Mean Number of Children Ever Born and Total Fertility Rate* by Religion and Religious Commitment, among Women 15 to 49 Years of Age, Puerto Rico, 1982

				Catholics	
	Total	Catholic	Non-Catholic	More Committed	Less Committed
Mean Number of Children Ever Born	1.8	1.8	1.8	1.8	1.9
No. of Cases	(3175)	(2144)	(1031)	(1204)	(940)
Mean Number of Children Ever Born to Women Ages 40–49	3.5	3.6	3.3	3.6	3.5
No. of Cases	(698)	(474)	(224)	(281)	(193)
Total Fertility Rate	2.5	2.5	2.5	2.2	2.8

* Calculated for the 5-year period prior to the survey, using cohort-period method.

The current total fertility rate is 2.5 for both Catholics and non-Catholics. Among Catholics, however, those who are more committed to their religion appear to have lower fertility than the less committed Catholics (2.2 vs 2.8), a finding the opposite of that expected.

Marriage/Nuptiality

The mean age at first marriage (Table 2) is essentially equal for Catholics and non-Catholics. This equality is also revealed in percent ever married by age, where little difference is observed between Catholics and non-Catholics. Among Catholics, however, religious commitment raises the age at marriage from 23.3 to 24.4 years compared to the low of 21.9 years for less committed Catholics. The difference in nuptiality between more committed and less committed Catholics is concentrated in the younger ages, with more committed Catholic women under the age of 30 showing 10 to 20 percentage points fewer marriages than less committed Catholics at the same ages.

Contraception

Current contraceptive use is essentially similar for Catholics and non-Catholics both in terms of overall levels and methods used (Table 3). Female sterilization is the most prevalent method used by both groups (39.6 and 39.8 percent). When religious commitment is considered, more committed Catholics show a higher total level of contraceptive use as well as higher method-specific rates for most of the methods compared to less committed Catholics.

When future birth intentions are taken into account (Tables 4 and 5), differences again are minimal. For women who have completed childbearing (Table 4), the total contraceptive prevalence as well as the method mix is very similar for Catholics and non-Catholics, with female sterilization by far the most prevalent method (around 57 percent). There also is little variation in contraceptive use between more and less committed members of the Catholic faith. The largest differences for women who intend to have no more children are observed in the sex-specific sterilization frequencies, with committed Catholic women having higher male sterilization (8.2 percent) and less committed Catholics having higher female sterilization (60.6 percent).

In Table 5, those women using contraception to space their births show only slight group differences, the most pronounced being a slightly greater preference for the pill among Catholic women compared to non-Catholics, with more committed Catholics less likely to use the pill and more likely to use the IUD, condom, and rhythm than less committed Catholics; all differences are small.

Breast-feeding

Breast-feeding differences are again minimal. The highest percentage of ever breast-feeding (33.2 percent) is among the less committed Catholics; the lowest is among non-Catholics (30.7 percent). Mean duration of breast-feeding is 4.4 months, varying from 4.8 months among less committed Catholics to 4.0 months among the more committed Catholics.

Discussion

Our results indicate that recent fertility of Catholics and non-Catholics in Puerto Rico is virtually identical and the

TABLE 2—Singulate Mean Age at Marriage (SMAM) and Percent Ever Married by Age, by Religion and Religious Commitment among Women 15–49 Years of Age, Puerto Rico, 1982

	Singulate Mean Age at Marriage						Catholics			
	Total		Catholic		Non-Catholic		More Committed		Less Committed	
SMAM (Years)	23.3		23.3		23.4		24.4		21.9	
	Percent Ever Married						Catholics			
Age Group	Total	No. of Cases	Catholic	No. of Cases	Non-Catholic	No. of Cases	More Committed	No. of Cases	Less Committed	No. of Cases
15–19	16.4	(609)	16.4	(404)	16.5	(205)	8.7	(261)	30.2	(143)
20–24	50.2	(527)	47.5	(368)	56.3	(159)	38.8	(179)	55.7	(189)
25–29	76.9	(482)	77.8	(313)	75.4	(169)	72.9	(151)	82.3	(162)
30–34	90.4	(438)	91.2	(295)	88.1	(143)	90.1	(154)	92.4	(141)
35–39	94.7	(421)	94.3	(290)	94.1	(131)	96.1	(178)	91.5	(112)
40–44	95.9	(384)	96.3	(267)	95.1	(117)	98.0	(157)	94.0	(110)
45–49	97.7	(314)	97.4	(207)	98.4	(107)	98.0	(124)	96.5	(83)

TABLE 3—Contraceptive Use by Religion and Religious Commitment, for Currently Married Women 15–49 Years of Age, Puerto Rico, 1982

				Catholics	
Contraceptive Use	Total	Catholic	Non-Catholic	More Committed	Less Committed
Currently Using	70.4	71.1	68.9	74.7	67.1
Female Sterilization	39.7	39.6	39.8	41.0	38.1
Male Sterilization	4.4	4.0	5.1	6.1	1.6
Pill	9.3	10.3	7.4	9.1	11.7
IUD	4.1	4.1	4.0	4.3	3.9
Condom	4.6	4.7	4.4	5.6	3.7
Rhythm	5.5	5.6	5.5	7.1	3.8
Other	2.8	2.8	2.7	1.5	4.3
Not Currently Using	29.6	28.9	31.1	25.3	32.9
Total	100.0	100.0	100.0	100.0	100.0
No. of Cases	(1557)	(1040)	(517)	(558)	(482)

TABLE 4—Contraceptive Use by Religion and Religious Commitment for Currently Married Women 15–49 Years of Age Who Intend No More Children, Puerto Rico, 1982

				Catholics	
Contraceptive Use	Total	Catholic	Non-Catholic	More Committed	Less Committed
Currently Using	82.5	83.4	81.1	84.2	82.0
Female Sterilization	57.3	57.5	56.9	55.1	60.6
Male Sterilization	6.3	5.8	7.3	8.2	2.6
Pill	5.6	6.0	4.8	6.1	5.9
IUD	2.9	2.7	3.4	2.6	2.8
Condom	3.8	4.2	3.0	4.9	3.3
Rhythm	4.4	4.7	3.9	6.3	2.5
Other	2.2	2.5	1.8	1.0	4.3
Not Currently Using	17.5	16.6	18.9	15.8	18.0
Total	100.0	100.0	100.0	100.0	100.0
No. of Cases	(1091)	(722)	(369)	(419)	(303)

importance of the proximate determinants through which women control their fertility is similar for Catholics and non-Catholics. Although we were unable to measure abortion, the similarity in levels of the three other proximate determinants and in total fertility, according to the Bongaarts model, implies little or no difference in abortion incidence across Catholic and non-Catholic categories.

Differences were observed between less committed and more committed Catholics in total fertility rate, age at first marriage, and contraceptive use, but these differences were not large and were in the opposite direction from the expected.

The finding of no effect of religious affiliation on fertility is consistent with earlier studies conducted in Puerto Rico. In a 1946 study, King claimed that although the Catholic Church had an intransigent position against all forms of contracep-

TABLE 5—Contraceptive Use by Religion and Religious Commitment, for Currently Married Women 15–49 Years of Age Who Intend More Children, Puerto Rico, 1982

Contraceptive Use	Total	Catholic	Non-Catholic	Catholics	
				More Committed	Less Committed
Currently Using	42.8	44.1	40.4	46.7	42.1
Female Sterilization	00.0	00.0	00.0	00.0	00.0
Male Sterilization	00.0	00.0	00.0	00.0	00.0
Pill	17.8	19.9	13.5	17.7	21.6
IUD	6.6	7.3	5.3	9.2	5.8
Condom	6.3	5.7	7.6	7.5	4.3
Rhythm	8.1	7.5	9.2	9.4	6.1
Other	4.0	3.7	4.8	2.9	4.3
Not Currently Using	57.2	55.9	59.6	53.3	57.9
Total	100.0	100.0	100.0	100.0	100.0
No. of Cases	(466)	(318)	(148)	(139)	(179)

tion, it did not appear that the Church was an important deterrent to the use of contraception by Puerto Ricans.[14,15] Ten years later, Stycos, *et al*,[16] discovered that despite over 80 percent of Puerto Ricans classifying themselves as Catholics, attitudes and behaviors were predominantly favorable to the small family achieved by the use of birth control.

Our findings also fit with results of another recent study that reveals widespread adoption of contraception among the Puerto Rican population. In particular, variables that would suggest different social environments, such as socioeconomic status or urban/rural residence, were found to have little or no effect on contraceptive use in this population.[17] It is very likely that the small geographic area of the island may produce homogeneity with respect to contraceptive knowledge and contribute to broad accessibility of contraceptive techniques.

On the other hand, socioeconomic status has been observed to be associated with marriage patterns in this population, with more highly educated women marrying at later ages than less educated women.[18] However, as the Catholic and non-Catholic groups in our study have similar distributions of educational level, it is unlikely that controlling this variable would have affected our analysis.

An important implication of the findings in this study is that Catholicism is not an obstacle to effective contraceptive use among Puerto Ricans. Thus, public health workers in family planning on the island of Puerto Rico and on the US mainland need not be unnecessarily distracted by the role of Catholicism in the adoption of effective contraception among Puerto Ricans and can focus their efforts on other family planning issues of concern to the Puerto Rican community.

What bearing does our study have on the Catholic fertility debate in the United States? First, the total fertility rate (TFR) of 2.5 in Puerto Rico in 1982 is higher by 0.7 child than in the United States (TFR = 1.8 in 1982).[7] Since over two-thirds of Puerto Ricans are Catholic, their migration to the United States would contribute to a higher aggregate fertility rate for Catholics compared to non-Catholics. In such a case, however, fertility differences would be the result of differences in the ethnic composition of Catholics and non-Catholics and not the result of religion *per se*.

Our findings support the *convergence* argument because we show no difference in the fertility behavior of Puerto Rican Catholics and non-Catholics. Our results are not consistent with the finding of Mosher, *et al*,[7] that fertility was higher for Catholic Hispanics than for non-Catholic Hispanics in the United States. The primary explanation for this discrepancy is that the effect of religion on fertility may differ by ethnicity among Hispanics. Puerto Ricans comprise a much smaller proportion of the US Hispanic population than Mexican Americans, who make up 60 percent of the Hispanic population and have substantially higher fertility than Puerto Ricans.[19,20] If Catholics have higher fertility than non-Catholics among Mexican Americans, this would affect the relation between religion and fertility for the total Hispanic population. Our discrepancy with the findings of Mosher, *et al*, point to the importance of disaggregated research on the Hispanic population if we are to resolve the debate over the effect of Catholicism on fertility in the United States.

APPENDIX

Religion

1.5 ¿ Es usted católica, protestante, de alguna otra religión, o no pertenece a ninguna religión?

1__ Católica
2__ Protestante
3__ Otro, especifique _____
4__ Ninguna

1.7 ¿ Cuán frecuentemente asiste usted a servicios religiosos?

1__ Nunca
2__ Una vez al año o menos
3__ Varias veces al año
4__ Una vez al mes
5__ Dos o tres veces al mes
6__ Una vez por semana
7__ Más de una vez por semana

Fertility

2.2 ¿ Cuántos bebés nacidos vivos ha tenido usted?

____
de nacimientos vivos

2.3 ¿ En qué fecha nació su (primer, segundo, etc.) hijo (a)?

	Primer Nacimiento Vivo	Segundo Nacimiento Vivo	Tercer Nacimiento Vivo	Cuarto Nacimiento Vivo	
Mes	____	____	____	____	Etc.
Año	____	____	____	____	

(ANOTE EN EL CALENDARIO COMO NV #1, etc.)

Religion

1.5 Are you Catholic, Protestant, some other religion, or no religion?
1__ Catholic
2__ Protestant
3__ Other, specify _____
4__ None

1.7 About how often do you usually attend religious services?
1__ Never
2__ Once a year or less
3__ Several times a year
4__ Once a month
5__ 2 or 3 times a month
6__ Once a week
7__ More than once a week

Fertility

2.2 How many live births have you had altogether?

____
live births

2.3 Thinking about your (1st/2nd/etc.) live birth, on what date was the baby born?

	First Live Birth	Second Live Birth	Third Live Birth	Fourth Live birth	
Month	____	____	____	____	Etc.
Year	____	____	____	____	

(RECORD ON CALENDAR AS NV #1, etc.)

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Nurse Midwifery Tuition Grant at University of Pennsylvania

The University of Pennsylvania Graduate Program in Nurse Midwifery has been awarded a grant from the US Department of Health and Human Services, Bureau of Maternal and Child Health and Resources Development, which provides tuition funds for a limited number of nurse midwifery trainees. Minimum criteria for application require a bachelor of science in nursing degree, leadership qualities, and a commitment to work in Title V or other publicly funded maternity care programs after graduation.

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Acculturation and Low Birthweight among Latinos in the Hispanic HANES

RICHARD SCRIBNER, MD, MPH, AND JAMES H. DWYER, PHD

Abstract: Self reports from 1,645 Latino mothers of Mexican descent who participated in the Hispanic Health and Nutrition Examination Survey (HHANES) were used to relate the birthweight of their infants to the HHANES acculturation index. After controlling for parity, a one point increase on the acculturation scale was found to be associated with a 1.19 (95% CI = 1.05, 1.34) increase in risk of maternal low birthweight (LBW) (1.98 risk increase for four points). The estimated relative risk increased to 1.34 (1.12, 1.60) with controls for age at interview, wealth, city size, and years of education; controlling for current smoking status reduced the relative risk to 1.31. US-born respondents were also at increased risk relative to Mexican-born, but this relation was explained by acculturation. The effect of education was found to depend on level of acculturation. Years of education was unrelated to risk among the Mexican-oriented, while increased education was associated with reduced risk in the US-oriented. These results suggest that factors associated with a Mexican cultural orientation may be protective against the risk of LBW. (*Am J Public Health* 1989; 79:1263–1267.)

Introduction

Concern about the relatively high rate of infant mortality in the United States has focused attention on the epidemiology of low birthweight (LBW), the major predictor of mortality in the first year of life.[1] In this regard, Latinos of Mexican descent present a surprising pattern. On the one hand they have one of the highest birthweight-specific fetal and neonatal mortality rates for LBW births of any ethnic group in the United States.[2] This finding is consistent with their status as a medically underserved minority,[3] and it may reflect a lack of medical intervention in the perinatal period.[4,5] On the other hand, the infant mortality rate for Latinos of Mexican descent is comparable to other Whites and is half that of Blacks.[6] This paradox has been explained by recent data on the distribution of LBW among Latinos.[2,6] Latinos of Mexican descent have one of the lowest risks of LBW births of any racial or ethnic group on which data are available. Thus, data on the prenatal experience of Latinos of Mexican descent may be informative concerning the conditions under which certain social and demographic risk factors lead to increased risk of LBW and infant mortality.[2]

In addition to a more favorable birth experience among Latinos relative to other US ethnicities and minorities, recent studies have indicated that nativity status (country of birth) of California Latino mothers is associated with the distribution of infant birthweights within the Latino population.[2] Furthermore, this association favors those mothers born in Mexico even though they are generally of lower socioeconomic status than those born in the United States.[6] This pattern of findings may be explained by at least two models: 1) Latino mothers who migrate are healthier than those who do not migrate; 2) birth in Mexico is merely a marker for a lifestyle that is protective against negative social influences on pregnancy outcomes. The purpose of this study is to explore the consistency of these competing explanations with data on Latino mothers from the Hispanic Health and Nutrition Examination Survey (HHANES).

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Methods

The HHANES was conducted by the National Center for Health Statistics during 1982–84.[7] Three separate Latino populations (Cuban, Mexican, and Puerto Rican) within the United States were sampled. Data for the current study came from the Mexican respondents included in the Medical History portion of the HHANES.

The degree to which respondents have taken on the social and cultural characteristics of either a Mexican or US orientation was assessed using an index of acculturation. The index consisted of eight variables selected from an index of 21 variables developed by Cuellar in 1980 for Mexican Americans.[8] These eight variables were measured by 18 survey items concerned with various aspects of acculturation. Four of the acculturation variables measure language preference, three measure ethnic identification, and one measures nativity status. The first seven variables were scored from 1 (Mexican-orientation) to 5 (US-orientation); the nativity variable ranged from 1 (respondent and parents born in Mexico) to 4 (respondent nor parents born in Mexico). The acculturation score is the arithmetic mean of the scores for the eight variables (i.e., total score/8). The resulting scale ranges from 1.0 (Mexican-orientation) to 4.9 (US-orientation). For purposes of the current study, this index was also disaggregated into a seven-item non-nativity acculturation index (ranging from 1.0 to 4.9) and a nativity index (ranging from 1 to 4).

The mother's status with regard to an LBW birth was assessed from three questionnaire items:
1) "What is the total number of live births you have had?"
2) "Did the (only) child weigh less than 5 1/2 pounds (2,500 grams) at birth?"
3) "How many of your children (who were born alive) weighed less than 5 1/2 pounds (2,500 grams) at birth?"

Female respondents were classified as mothers if they indicated one or more live births in the first item. A mother was then classified as a LBW mother if she indicated one or more LBW births in either of the latter two questions (no = 0, yes = 1).

Smoking status was determined from the questions: "Do you currently smoke cigarettes?" Level of education was measured by reported years of education. Wealth index was measured by the ratio of average household income for the past 12 months over the poverty threshold income for 1982. Scores of less than 100 were below the poverty level for 1982. City size was an ordinal variable with 8 categories ranging from 1 (cities over 1,000,000) to 8 (towns less than 10,000).

The data were analyzed using logistic regression because the dependent variable was dichotomous. The model used was the following:[9]

$$\ln(\text{odds}[\text{Maternal LBW}]) = \alpha\ P1 + \beta_1\ P2 \ldots + \beta_x\ ACC + \beta_{x+1}\ AGEI$$

Where maternal LBW, parity of one (P1) and parity of two (P2) are dichotomous variables, and acculturation score (ACC) and age at interview (AGEI) are continuous variables. The CATMOD procedure in the software package SAS was used to analyze the data. Survey weights were incorporated by dividing each individual weight by the mean survey weight of all the Latino mothers analyzed yielding an average sample weight of one. Design effects were not incorporated in the analysis. The degrees of freedom for estimating the design effects in Hispanic HANES are small and therefore yield unstable point estimates. Test statistics generated using unstable point estimates of the variance could be misleading. However, the use of logistic regression for the analysis causes the design effect to converge on 1.0. In addition, including variables related to survey design (e.g. age, sex) also causes the design effect to converge on 1.0.*

Results

Total births ranged from one to 15. Mothers were eliminated from the analysis if data were unavailable on total births, acculturation index, or number of LBW births. One mother reported 11 births of which 11 were LBW births. The next highest number of reported LBW births was 4. The mother with 11 low birthweight births was 58 years old, born in the US, had an acculturation score of 2.0, and a survey weight above the average. She was the only case deleted from the analysis.

The 1,645 mothers reported a total of 6,216 births. Of these births, 299 were reported as LBW births. The overall proportion of LBW infants (4.8 percent) was comparable to the proportion of LBW infants for Mexican-Americans found by Williams[2] (4.7 percent) and Ventura[6] (5.3 percent) in cross-sectional analyses of California state and national vital records, respectively, in 1981. The percentage of LBW infants among Mexican born respondents in the HHANES (3.9 percent) was also comparable with that found by Williams (4.2 percent), as were the proportions of LBW infants among United States born respondents (5.5 percent) when compared with Williams (5.2 percent). However, the proportions of LBW infants among US-born (6.3 percent) and Mexican-born (5.0 percent) mothers reported by Ventura were higher than those found both by Williams and in the HHANES.

The distribution of the mothers on the acculturation index was bi-modal which may reflect the presence of two fairly distinct subpopulations of Latinos of Mexican descent (Figure 1). To explore the characteristics of these two subgroups defined by this distribution the mothers were dichotomized, those equal to or below an acculturation score of 2.4 were designated as Mexican-oriented and those above 2.4 designated as US-oriented (Figure 1). Demographically, Mexican-oriented mothers had a higher mean parity and lower means for years of education and income. Behaviorally, Mexican-oriented mothers were less likely to report current smoking of cigarettes (Table 1).

*Personal communication with Lester Curtain of the National Center for Health Statistics.

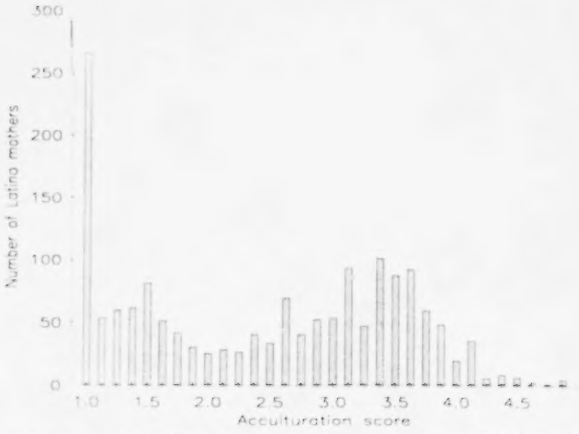

FIGURE 1—Distribution of all Latino Mothers in Hispanic HANES by Acculturation Score (n = 1.645)

TABLE 1—Variable Means by Cultural Orientation

Demographic/ Behavioral Variables	Mexican-Orientation			US-Orientation		
	n	Mean	S.D.	n	Mean	S.D.
Years of education	731	6.05	3.82	914	10.27	3.29
Wealth index (<100 poverty)	661	138.8	96.6	840	191.5	135.6
Age (years) at interview	731	42.1	14.3	914	38.5	14.0
City size (1 = urban, 8 = rural)	680	3.75	2.24	824	5.22	2.35
Parity (# live births)	731	4.24	2.95	914	3.41	2.44
Mexican-born (0 = no, 1 = yes)	731	.735	.447	914	.130	.334
Current Smoker (0 = no, 1 = yes)	731	.194	.396	914	.281	.450

The percentage of LBW mothers in the survey was 13 percent (11 percent among the Mexican-oriented women, 14 percent among US-oriented women). Mexican-oriented women were at lower risk of maternal LBW status within all parity groups except after primiparity (Figure 2).

When LBW status of mothers was logistically regressed on the acculturation score the coefficient for acculturation was positive ($\hat{\beta}$ = .080, S.E. = .067) but close to zero (Table 2). When parity was added to the model, the acculturation coefficient increased in magnitude ($\hat{\beta}$ = .172, S.E. = .071), reflecting the tendency of greater parity among Mexican-oriented women to mask the relation between LBW risk and acculturation. The regression coefficients for demographic variables (education, wealth index, age at interview, and city size) were all near zero with parity controlled.** However, current smoking status had a strong positive effect ($\hat{\beta}$ = .355 S.E. = .160) and remained so ($\hat{\beta}$ = .396, S.E. = .168) with controls for acculturation score. Notably, the positive effect of acculturation score on maternal LBW status was reduced ($\hat{\beta}$ = .157, S.E. = .071) after controlling for current smoking status (Table 2).

Education was the only variable that had a substantial

**Data available on request to author.

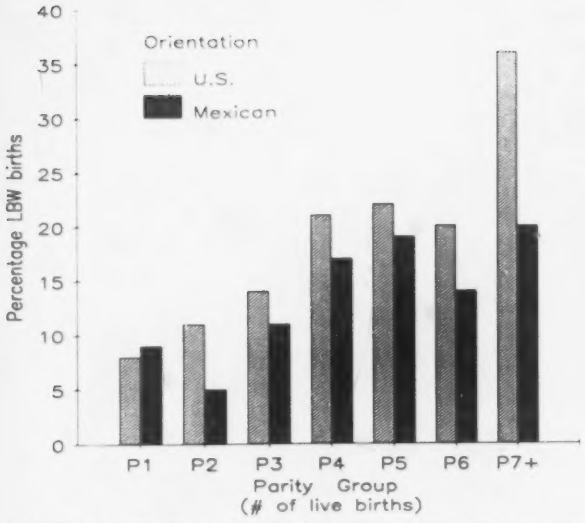

FIGURE 2—Percentage of Mothers with Low Birthweight Births for each Cultural Orientation by Parity Group

interaction with acculturation (Table 3). The interaction effect was negative ($\hat{\beta}$ = −.063, S.E. = .021) as was the main effect of education ($\hat{\beta}$ = −.069, S.E. = .026) while the main effect of acculturation ($\hat{\beta}$ = .234, S.E. = .081) was positive. This indicates that the effect of years of education on maternal LBW status varies with acculturation status. Specifically, years of education has no effect on maternal LBW status for Mexican-oriented women but a negative effect among US-oriented women (Figure 3).

To explore the independent effect of nativity, this component of the acculturation index was removed and maternal LBW status was logistically regressed on nativity and parity. The effect of nativity on risk was positive ($\hat{\beta}$ = .116, S.E. = .055), and increased ($\hat{\beta}$ = .158, S.E. = .068) with controls for demographic variables. The acculturation index with the nativity variable removed (i.e., a non-nativity based acculturation index) was also positive and substantial both controlling for parity ($\hat{\beta}$ = .196, S.E. = .082) and controlling for parity and the other demographic variables ($\hat{\beta}$ = .311, S.E. = .113). When both the nativity variable and the non-nativity acculturation variable were included in the analysis the effect of nativity was reduced to near zero ($\hat{\beta}$ = .030, S.E. = .089) while non-nativity acculturation status remained correlated with maternal LBW status ($\hat{\beta}$ = .160, S.E. = .132). This relation was more pronounced when demographic and behavioral variables were added to the model. The effect for nativity remained small ($\hat{\beta}$ = .043, S.E. = .100) while the effect for non-nativity based acculturation increased ($\hat{\beta}$ = .259, S.E. = .165) (Table 4).

For ease of interpretation and additional assurance that the results are not the result of model specification all independent variables were dichotomized and reentered into the regression equation. The model including acculturation score and parity indicated that the risk of maternal LBW among US-oriented Latinos was 1.64 times that of the Mexican-oriented mothers. This relative risk increased to 1.86 when all covariates were included in the model.

Discussion

This study confirms the previously reported reduced risk of LBW among the Latino mothers born in Mexico relative to Latino mothers of Mexican-descent born in the United States.[2,6] However, this association was weaker than that between measured acculturation and maternal LBW risk, and became negligible when the non-nativity acculturation variable was included in the regression model.

The self-reporting of pregnancy outcome involves recall error. The recall period in the HHANES is as long as 50 years, and no queries were made concerning whether or not an accurate birthweight was supplied to the mother at the time of delivery. As a result, overestimating birthweight by Mexican-oriented women could account for these results. However, a similar reporting bias was felt to be responsible for the low perinatal mortality rates for Latino infants of Mexican-descent when the data from linked birth and death certificates first became available.[11,12] It was subsequently shown by Williams in California that these low rates could not be the result of a systematic bias from underreporting by Mexican mothers.[2] It is therefore noteworthy that in the HHANES the proportion of mothers' self-reports of low birthweight, both overall and among the Mexican- and US-born subgroups, were not substantially different than that

TABLE 2—Logistic Regression Coefficients and Standard Errors for each Numbered Model by the Independent Demographic Variables Included in the Model

Model	(1) n = 1645	(2) n = 1645	(3) n = 1645	(4) n = 1630	(5) n = 1501	(6) n = 1504	(7) n = 1645	(8) n = 1363	(9) n = 1363
Demographic Variables									
Acculturation	.080(.067)	.172(.071)	.177(.072)	.230(.084)	.182(.076)	.176(.075)	.157(.071)	.252(.092)	.228(.093)
Age at interview	.	.	.003(.006)	.	.	.	.	.004(.007)	.003(.007)
Years of education	.	.	.	−.031(.023)	.	.	.	−.021(.026)	−.020(.026)
Wealth index	.	.	.	.	−.001(.001)	.	.	−.001(.001)	−.001(.001)
City size	.	.	.	.	.	.005(.030)	.	−.029(.031)	−.027(.031)
Current smoker status	.	.	.	.	.	.	.355(.160)		.396(.168)
Parity = 1	.	*	−2.77(.349)	*	*	*	*	*	*
Parity = 2	.	*	−.303(.248)	*	*	*	*	*	*
Parity = 3	.	*	.276(.242)	*	*	*	*	*	*
Parity = 4	.	*	.775(.253)	*	*	*	*	*	*
Parity = 5	.	*	.762(.294)	*	*	*	*	*	*
Parity = 6	.	*	.477(.373)	*	*	*	*	*	*
Parity = 7+	.	*	1.16(.299)	*	*	*	*	*	*

*Parity was controlled with seven dichotomous dummy variables. Model 3 is demonstrating all the coefficients for that model.
Note: All the coefficients included in a model are shown in the column below the model number except as noted. The independent variable is maternal LBW status.

TABLE 3—Logistic Regression Coefficients and Standard Errors for each Numbered Model by the Independent Variables and Interaction Variable

	Model				
	(1)	(2)	(3)	(4)	(5)
Interaction[Φ]	−.063(.021)	.007(.005)	.000(.001)	.012(.031)	−.058(.079)
Acculturation score − 3	.234(.081)	.179(.071)	.188(.078)	.230(.080)	.184(.081)
Years of education − 8 years	−.069(.026)	.	.	.	.
Age of interview −39 years	.	.008(.007)	.	.	.
Wealth index − 170 score	.	.	−.001(.001)	.	.
City size − 4.9 score	.	.	.	−.004(.036)	.
Current smoker status	.	.	.	.	.388(.158)
Parity	*	*	*	*	*

ΦInteraction of acculturation and the other independent variable included in the model with maternal LBW status.
*Parity was controlled using seven dichotomous dummy variables.
NOTE: All the coefficients included in a model are shown in the column below the model number except as noted. The independent variable is maternal LBW status.

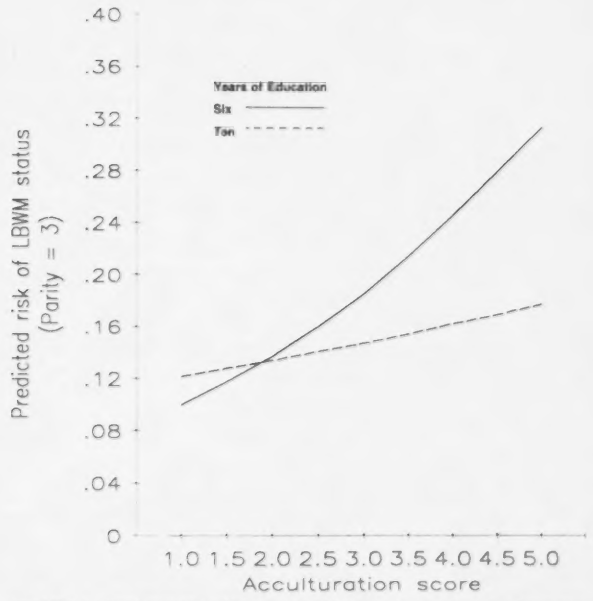

FIGURE 3—Predicted Risk of Maternal LBW Status by Acculturation Score and Years of Formal Education Demonstrating the Interaction of Acculturation and Education
Model: ln (Odds{Maternal LBW}) = $\alpha + \beta_1*(P3) + \beta_2*(ACC-3) + \beta_3*(EDUC-8) + \beta_4*((ACC-3)*(EDUC-8))$

found by Williams in California. Furthermore, numerous studies have found that birthweight is recalled with a high level of accuracy by mothers many years after the birth.[13–16]

Controlling for demographic risk factors in the regression model did not change the magnitude of effect of acculturation on maternal LBW status, in fact the effect became stronger. It may be that acculturation status acts as a marker for certain values, beliefs or lifestyles which protect these women from negative birth outcomes associated with the demographic pattern they possess.[2] Where formal education has little effect on maternal LBW status in Mexican-oriented mothers, it is negatively associated with maternal LBW status in US-oriented mothers. This observation is consistent with previous studies.[6,17] A study of infant mortality in Israel found that level of formal education in an out-group was directly related to infant mortality contrasted with the inverse relationship in the general population.[18]

Behaviorally, Latinos of Mexican descent demonstrate better nutrition, fewer premarital births, lower rates of smoking and alcohol use, and higher regard for parental roles when compared to other Whites.[2,19–22] Consistent with such findings, a higher proportion of the US-oriented Latino mothers responded that they currently smoked. Cigarette smoking was the only independent variable other than parity that has a strong effect on maternal LBW status both before and after controls for acculturation were added to the model. The effect of acculturation status, on the other hand, was reduced slightly by controlling for current smoking status, a behavior associated with the US orientation. The fact that the designation is an assessment of smoking status at the time of survey and not an indication of prenatal smoking probably

TABLE 4—Logistic Regression Coefficients and Standard Errors for each Numbered Model by the Independent Nativity Variables Included in the Model

Model	(1) n = 1645	(2) n = 1645	(3) n = 1645	(4) n = 1363	(5) n = 1424	(6) n = 1424
Nativity Variables						
Non-Nativity based acculturation score	.196(.082)	.	.160(.132)	.311(.113)	.	.259(.165)
Nativity (1 = Mexico, 4 = United States)	.	.116(.055)	.030(.089)	.	.158(.068)	.043(.100)
Age at interview	.	.	.	.003(.007)	.003(.007)	.003(.007)
Education	.	.	.	−.015(.027)	.004(.025)	−.013(.027)
Wealth index	.	.	.	−.001(.001)	−.001(.001)	−.001(.001)
City size	.	.	.	−.025(.035)	−.024(.036)	−.028(.036)
Current smoker status	.	.	.	.376(.175)	.396(.174)	.396(.174)
Parity	*	*	*	*	*	*

*Parity is controlled with seven dichotomous dummy variables.
NOTE: All the coefficients included in a model are shown in the column below the model number except as noted. The independent variable is maternal LBW status.

attenuates the association. Also, since no other behavior variables were controlled in the analysis, current smoker status could be a marker of a group of behaviors related to smoking and positively associated with maternal LBW status. This situation would exaggerate the apparent effect of this variable in the analysis.

Lack of prenatal care is often cited as a risk factor for LBW. However, Mexican-born Latinos have a lower rate of LBW than their US-born counterparts despite the fact they are more likely to receive late or no prenatal care.[6] These findings, together with those reported here suggest that Mexican-orientation is associated with behaviors or lifestyle predating pregnancy which are more important determinants of LBW than prenatal care. Thus utilization of prenatal care services among Latinos of Mexican descent may be associated with a lifestyle that places them at risk of LBW. Utilization of prenatal care may become more likely with the loss of Mexican values and beliefs which are protective against LBW.

These analyses demonstrate that Latinos of Mexican descent are not a homogeneous group with regard to risk for LBW. Furthermore, Mexican cultural orientation acts as a marker for a lifestyle which is associated with a favorable prenatal experience among these Latino mothers. Such an effect is seen primarily in the highly Mexican-oriented Latinos and lost when one moves away from this cultural orientation. It is noteworthy that conventional demographic risk factors for LBW, such as formal U.S. education, only become important for Latinos of Mexican descent as they move away from the Mexican orientation. This may be due to the lack of influence of these factors in the social environment that developed the culture which gave rise to the values and beliefs making up the Mexican orientation.

The components of the Mexican orientation which may provide protection are important from a prevention standpoint. Especially in light of the observation that this orientation may ameliorate the impact of demographic risk factors. Some possible components of the Mexican orientation that need further investigation include nutrition, familial integration, smoking and alcohol use, and out-of-wedlock births. How Latinos lose those attributes which proscribe adverse health behaviors and/or how they develop attributes which do not effectively proscribe them may be critical to our understanding of LBW.

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Fetal Death Ratios in a Prospective Study Compared to State Fetal Death Certificate Reporting

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Abstract: A cohort of 6,254 pregnancies surviving at least 20 weeks of gestation was identified through pregnancy testing and follow-up at three Kaiser Permanente medical offices in northern California in 1981–82. Fetal death ratios per 1,000 live births were 12.1 for all fetal deaths versus 5.0 for the subset of fetal deaths reported to the California state registrar Only fetal deaths resulting in overnight hospitalization of the mother were reported. Seventy-nine percent of fetal deaths over 28 completed weeks since the last menstrual period (LMP) were reported versus only 10 percent between 20 and 28 completed weeks since the LMP. Ninety-three percent of fetuses over 400 grams were reported. The unreported fetal deaths were mainly those perceived by attending physicians as spontaneous abortion, especially missed or incomplete spontaneous abortion. Physicians apparently preferred the label of spontaneous abortion over stillbirth or fetal death whenever fetal maturity could not be substantiated, regardless of prior estimates of the date of the LMP. Fetuses as large and developed as potentially viable infants were the most likely to be reported. (*Am J Public Health* 1989; 79: 1268–1270.)

Introduction

State fetal death certificate reporting has been the primary source for determining the risk of stillbirth, or late fetal death, in the United States for more than 60 years.[1] While the legal definition of reportable fetal death has varied somewhat by time and state, reasonably consistent trends in fetal death ratios (the number of fetal deaths per 1,000 live births) have been observed across the country and over the years. Nevertheless, underreporting of fetal deaths is a well-recognized and pervasive problem[2–4] whose magnitude is unknown.

Most states now require that a fetal death certificate be filed for spontaneous fetal deaths over 20 weeks of gestation. Eight states also require filing certificates for earlier fetal deaths, and four states specify fetal delivery weight over 350–500 grams as an alternative reporting criterion to 20 weeks gestational age.* As recently demonstrated by Harter, *et al*,[4] and others in the past,[2,3] states with reporting requirements based on gestational ages less than 20 weeks or based on fetal delivery weight have consistently higher ascertainment of fetal deaths between 20 and 28 weeks. After 28 weeks of gestation, reporting of fetal deaths appears to be fairly complete.

Harter, *et al*, imply that reasons for underreporting may be related to the condition of the fetus upon expulsion. If it is macerated, malformed, or very small, its chances of being reported are diminished. If the fetus dies in utero before 20 weeks of gestation but is carried for days or even weeks afterward, as in the case of "missed abortion," whether or not it should be reported may be unclear to physicians.

The same medical personnel responsible for reporting fetal deaths often use the International Classification of Disease (ICD) index to code hospital discharge summaries. A late fetal death (ICD-9 code 656.4) is defined in the ICD-9 Tabular List as having more than 22 completed weeks of gestation while spontaneous abortion appears with a separate code (634) presumably for earlier fetal deaths occurring prior to 22 weeks. This may also cause confusion.

In California, Section 10175 of the California Health and Safety Code, first prepared in 1955, states, "Each fetal death in which the fetus has advanced to or beyond the twentieth week of uterogestation shall be registered." An advisory was later circulated to county health departments in 1978 stating that if the gestational age cannot be determined, a weight of over 400 grams and a crown-heel length of over 28 centimeters may be used as alternative reporting criteria. While this advisory is still in effect, existing copies are difficult to find. I was unable to locate written documentation from the state vital registration office, the county offices, or Kaiser hospitals.

To see which fetal deaths are likely to be reported and which are not, I used data from the Kaiser Permanente Medical Care Program in northern California collected for a large prospective study of pregnancy outcome. In this study there were 75 fetal deaths of 20 or more weeks of gestation within a cohort of 6,254 pregnancies surviving at least 20 completed weeks from the last menstrual period (LMP). Gestation was counted from the first day of the LMP until expulsion from the womb. Days 0–6 constitute the 0th week so that days 140–146 constitute the 20th week (this convention was established in prior prospective studies of pregnancy outcome).[5–8]

Methods

The Kaiser program in northern California is one of the oldest health maintenance organizations in the country, now providing prepaid medical coverage to about 2 million members. Ninety percent of the membership join through employee organizations and 10 percent as individuals. Kaiser members tend to report early for medical care, including pregnancy diagnosis; they also tend to stay with the program for long periods of time.

In early 1983, researchers attempted to identify all women having positive pregnancy tests or pregnancy confirmation at three Kaiser clinics in the south San Francisco Bay Area, September 15, 1981 through June 30, 1982. The purpose was to establish a cohort of women in early stages of pregnancy during the nearly one-year period that the pesticide, malathion, had been sprayed in the area by the State of California Department of Agriculture. A case-control study was then embedded within the cohort to study the effects of

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the spraying and of other environmental exposures during pregnancy.[9–11]

An original cohort of 9,564 pregnancies was identified by scanning logs of urine pregnancy tests and pregnancy confirmation visits at the three clinics (Table 1). Pregnancy outcomes were determined from computer linkage to hospital discharge summaries, California birth certificate files, and Kaiser abortion clinic data. For outcomes not obtained from these three sources, individual medical records were reviewed. Records were then re-reviewed for confirmation of all outcomes other than live birth indicated in Table 1. Particular attention was given to determining as accurately as possible the gestation period for each pregnancy.

Of the 9,564 pregnancies, 6,254 survived at least 20 completed weeks and remained under care at one of the three clinics through the time of delivery. Of these, 75 ended in spontaneous fetal death and 6,179 ended in live birth. These form the basis of this study on reporting practices for fetal deaths. Descriptions of the pregnancies ending in fetal death were derived both from medical records at the three clinics and from the program's computerized hospital discharge summaries.

An extensive search was made of state fetal death certificate files to determine whether the fetal deaths had been reported to the state registrar. The search was conducted by manually matching the study subjects to four separate lists of fetal death certificate files sorted by hospital, by mother's last name, by mother's first name, and by date, respectively.

Results

Forty-one of the 75 fetal deaths occurred before 28 completed weeks of gestation (between day 140 and day 195, inclusive) and 34 occurred after 28 completed weeks (day 196 and beyond). Most (37/41) of the 20–27 week fetal deaths were not reported, while most (27/34) of the 28–41 week fetal deaths were reported. Only about half (21/41) of the 20–27 week fetal deaths resulted in an overnight hospitalization of the mother, while most (30/34) of the 28–41 week fetal deaths resulted in overnight hospitalization of the mother. For the 24 fetal deaths not requiring overnight hospitalization all were attended by physicians: nine in the emergency rooms and 15 only in the clinics. Dilation and curettage procedures were performed on 16 of the 24.

Table 2 shows the diagnosis codes for fetal outcome on the mother's computerized hospital discharge summary by fetal death certificate reporting status. Three of the 31 mothers with reported fetal death showed no code for fetal death, or for abortion—only miscellaneous conditions of the delivery such as preeclampsia or premature rupture of membranes. A search for fetal deaths using hospital discharge diagnoses only would miss these cases. For the 20 mothers whose fetal deaths were not reported even though hospitalization had occurred, most had codes for some kind of abortion (spontaneous, missed, elective, molar pregnancy or abortion type unspecified). Four had the appropriate ICD-9 code 656.4 for intrauterine death, three of whom delivered fetuses under 400 grams. Of the 24 other women who were not hospitalized, none of their fetal deaths were reported to the state.

The 31 reported fetal deaths fell into several categories: cord problems, placental problems, abnormal presentation, twin pregnancy, sepsis, complications of toxemia, congenital anomaly, and unknown cause of death. (In only four cases was the cause of death unknown by the attending physician.) None of the reported fetal deaths were described in the medical chart as missed abortion. Birthweights were generally considerably over 400 grams although three were less than or equal to 400 grams. Four of the reported fetal deaths were less than 28 weeks gestation (three at 26 and one at 24 weeks); the remainder were all over 28 completed weeks.

Of the 37 unreported fetal deaths occurring during the 20th through 27th weeks, 28 (76 percent) were referred to in the medical chart as missed or incomplete abortion, one was a molar pregnancy, and the remaining eight were described as complete spontaneous abortions. Delivery weights were generally unknown or not recorded. Two delivery weights, however, were over 400 grams. Both resulted in hospitalization of the mother but with a spontaneous abortion code rather than the intrauterine death code on the hospital discharge summary. Of the seven unreported fetal deaths over 28 completed weeks of gestation, only one was over 400 grams. All were under 34 weeks at the time of expulsion. Five of the seven were referred to in the medical chart as missed abortions with the fetus described as autolyzed, mummified, or macerated. Causes of death were unknown or unspecified in over half of the unreported fetal deaths. Known causes of death, however, were similar to the reported fetal deaths.

TABLE 1—Fetal Outcomes in Cohort and in Fetal Death Study among 9,564 Pregnancies Identified at Three Kaiser Medical Offices, September 15, 1981–June 30, 1982

Fetal Outcome	Number In Cohort	Number in Fetal Death Study[c]
Live Birth	6,759	6,179
Fetal Death ≥20 Weeks	80	75
Fetal Death <20 Weeks	538	—
Induced Abortion for Malformation	5	—
Voluntary Induced Abortion	1,802	—
Ectopic Pregnancy	48	—
Unknown	391[a]	—
Total	9623[b]	6,254

[a]Unknowns were largely due to terminated membership in the Kaiser program.
[b]Includes 114 co-twins and three co-triplets.
[c]Includes all that survived at least 20 weeks and remained under care at one of the three Kaiser facilities.

TABLE 2—Diagnosis of Fetal Outcome on Mother's Hospital Discharge Summary by Fetal Death Certificate Reporting Status

Diagnosis	ICD-9 Code	Reported	Not Reported
Intrauterine death ≥22 weeks	656.4	28	4
Abortion			
Spontaneous	634	—	2
Missed	632	—	8
Elective	635	—	2
Molar pregnancy	630	—	1
Unspecified	637	—	2
Miscellaneous condition of delivery[a]	—	3	1
Not hospitalized	—	0	24
Total	—	31	44

[a]No mention of fetal outcome.

Fetal death ratios (the number of fetal deaths divided by the number of live births per 1,000) were calculated for reported fetal deaths, for hospitalized fetal deaths, and for all fetal deaths (Table 3). For fetal deaths at greater than 28 completed weeks of gestation, there were only small differences between the three nested categories. For fetal deaths defined from the 20th week, the fetal death ratio for all fetal deaths was more than double that for only reported fetal deaths; the ratio for hospitalized fetal deaths was midway between the other two categories.

Discussion

These findings are consistent with other prospective investigations conducted in the United States with follow-up through full term[5–7]: fetal death ratios for fetal deaths beyond 20 weeks of gestation were 40–80 percent higher than the respective state averages. In another study in which follow-up was only through the end of the 27th week,[8] fetal death rates for weeks 20–27 were consistent with the present study.

While death in utero after 20 weeks of gestation seems a simple definition, persons responsible for reporting fetal deaths are either confused by the definition or ignore the requirement. Factors other than time since LMP appear to play important roles in reporting fetal deaths. In the Kaiser data, overnight hospitalization must have occurred for a fetal death to be reported. But even with hospitalization, only 68 percent of fetal deaths over 20 completed weeks of gestation were reported to the state. In contrast, 89 percent of hospitalized fetal deaths over 28 completed weeks were reported. The best predictor of reporting was birthweight, with 93 percent of weighed fetuses delivered over 400 grams being reported.

Reporting practices found in the Kaiser program may be representative of reporting practices elsewhere. Fetal death ratios at Kaiser hospitals based on reported fetal deaths are about average in the state of California.[12] Although guidelines[13] exist to help clarify the difference between fetal death and neonatal death there are no routinely distributed guidelines from the states and counties to help distinguish reportable fetal death from other fetal deaths.

Perhaps physicians may rely more on their obstetric training than on state regulations to determine reportability. *Williams Obstetrics*,[14] a major textbook for obstetric training in this country, emphasizes birthweight as the main criterion for reportability, stating ". . . fetuses weighing less than 500 grams usually are not considered as births, but rather as abortions, for purposes of perinatal statistics." Gestational age is then referred to in terms of equivalency to birthweight.

It is routine practice at Kaiser to offer a woman options for disposition of the body of the fetus if it weighs over 500 grams. Such options are not routinely offered for smaller fetuses.

If the trends in this study are representative of widespread practices, the majority of underreporting of fetal deaths may be due not to negligence or random omissions, but rather to systematic perception of which fetal deaths are reportable. Fetuses under 500 grams may seldomly get reported regardless of time since LMP.

This practice in fetal death reporting may account for the consistently lower rates of fetal death found in studies based on state vital statistics compared to prospective studies based on follow-up of pregnancy cohorts. In the prospective studies, gestational age at outcome is based on time between LMP and expulsion, not on birthweight equivalencies.

TABLE 3—Fetal Death Ratios per 1,000 Live Births

	≥20 Completed Weeks of Gestation	≥28 Completed Weeks of Gestation
Fetal death certificate	5.0	4.4
Hospitalized	8.3	4.9
All fetal deaths	12.1	5.5

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Effects of a Geriatric Nurse Practitioner on Process and Outcome of Nursing Home Care

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Abstract: We compared measures of quality of care and health services utilization in 30 nursing homes employing geriatric nurse practitioners with those in 30 matched control homes. Information for this analysis came from reviews of samples of patient records drawn at comparable periods before and after the geriatric NPs were employed. The measures of geriatric nurse practitioner impact were based on comparisons of changes from pre-NP to post-NP periods. Separate analyses were done for newly admitted and long-stay residents; a subgroup of homes judged to be best case examples was analyzed separately as well as the whole sample. Favorable changes were seen in two out of eight activity of daily living (ADL) measures; five of 18 nursing therapies; two of six drug therapies; six of eight tracers. There was some reduction in hospital admissions and total days in geriatric NP homes. Overall measures of medical attention showed a mixed pattern with some evidence of geriatric NP care substituted for physician care. These findings suggest that the geriatric NP has a useful role in nursing home care. (*Am J Public Health* 1989; 79:1271–1277.)

Introduction

Nursing home care has been the subject of widespread concern and criticism. Within professional circles, it enjoys little prestige. It is viewed as low technology care with little hope of patient improvement and little opportunity to use professional skills.[1] Physician visits tend to be infrequent, conforming to the minimal levels required to authorize the care of residents.[2]

Ironically, when care is provided to nursing home residents they show great responsiveness. In fact, almost any intervention produces a positive response.[3] One promising approach to improve the care of the elderly in nursing homes engages nurses to work as geriatric nurse practitioners (GNPs) to provide or coordinate primary care for residents. An early study in Salt Lake City showed that the geriatric NPs could work effectively as part of a multi-disciplinary team to improve the function of nursing home residents and reduce the costs of hospital care.[4] A study in Boston showed similar improved care and savings were possible.[5]

The enthusiasm of these results has not been matched by patterns of funding. Direct payment for the services of geriatric NPs was not allowed under Medicare Part B if the services were given without on-site physician supervision. Thus geriatric nurse practitioner care in nursing homes was discouraged.

With the support of the W.K. Kellogg Foundation, beginning in 1976, the Mountain States Health Corporation [MSHC] arranged for nurses working in nursing homes throughout four northwestern states to be trained through the continuing education (CE) programs at three nursing schools. The training used the already established CE model of four months of didactic training at the university and eight months of preceptorship with a local physician at the home site. The nursing homes agreed to add the geriatric NPs to their staffs and to allow them to function in the expanded role for which they had been trained. A contract for a minimum of 18 months of employment after training was negotiated, with the geriatric nurse practitioner being an employee of the nursing home rather than an independent contractor or an employee of a physician group. In 1982, the program expanded to cover 13 western states, using four western schools of nursing to provide the continuing education.

After the program was well established, an evaluation of the effect of adding these GNPs on nursing home costs and quality of care was undertaken. The study design had three major components: a prospective study of the functional changes seen in a sample of residents treated by geriatric NPs compared with matched controls, a retrospective review of the records from the geriatric nurse practitioner homes and matched controls, and an analysis of the costs to both payors and nursing homes associated with this innovative approach to care.[6] This paper describes the results of the retrospective study.

Methods

Because the study was mounted after the program was well established, there was no opportunity to design a randomized clinical trial. From the pool of about 100 nursing homes then employing trained geriatric nurse practitioners, 30 pairs were developed by matching each GNP nursing home to a control home in the same state* on the basis of proportion of Medicare cases, ownership, corporate management, number of beds, and rural/urban location. Each potential match required both a GNP home and a control home willing to allow access to records (and for some, patients).

A quasi-experimental design was used, consisting of pre- and post-geriatric nurse practitioner periods for both groups of homes. The data source was the resident's nursing home records. The pre-GNP period consisted of one year prior to the geriatric NPs' employment in that role, and the post-GNP period began *with* the NP's employment and lasted up to two years for any resident (or until the resident was permanently discharged from the home). The training period was not part

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of the study. The time periods for the control home coincided with those for the paired geriatric nurse practitioner home. Once these general time periods were established for each sample pair, some minor adjustments were made (never more than a six-month shift) to accommodate the dates of the study to the home's fiscal year in order to link financial and practice data. This time-based sampling allowed a cross-section of periods among the 30 pairs that encompassed more than eight years of practice and changes in the external environment.

Within each nursing home, two samples of residents' records were selected randomly—one for the pre-period and a second for the post-period. The same number of records was sought in each home, resulting in varying sampling ratios depending on numbers of beds and turnover rates. The majority of the homes (or the wings that used a geriatric nurse practitioner) were approximately 100 beds. Within each sample, a 2:1 ratio of newly admitted to longer stay residents was planned; however, it was not possible to achieve this ratio in some cases because of insufficient numbers of newly admitted cases. As shown in Table 1, the final pre-GNP sample was more equally weighted. Because the post-GNP period was two years, a slight variation in the sampling method was used. The sample was divided into the two years but the second year sample added only new admissions.

Demographic information, functional status, nursing therapies, medications, medical attention, and use of hospital and emergency room services were abstracted from the medical records. The general time frame used for this component was to assess the resident's condition on entrance to the nursing home (or the beginning of the study period for long stays) and at discharge (or end of study period). For selected variables, a third status point was used at three months, the median stay for most admissions. For many variables where use varied with time, a two-week window was defined and the rates calculated for that period.

In addition to tabulating numbers of medications given (separated by regular and PRN (as occasion requires) use), specific dose-equivalents of six types of drugs most commonly used in nursing homes were calculated: psychotropics, sedatives, tricyclic antidepressants, diuretics (except furosemide), furosemide, and digoxin.

Orders for various services (medications, laboratory tests, nursing orders, and special services) were tabulated according to the person ordering them and the mode of ordering them (i.e., in person or by telephone). Visits by attending physicians and specialized personnel (podiatrists, dentists, physical therapists, occupational therapists) were tabulated. The frequency of services was calculated as a rate per day of stay in the nursing home. Use of emergency room services was separated by whether the visit was prompted by an emergency or for routine testing. Hospital admissions and lengths of stay were recorded. The data on hospital use are again presented as a rate per day, using the patient's total length of stay in the nursing home as the denominator.

A series of tracers was developed for conditions for which record data might be available (diabetes, congestive heart failure, hypertension, new urinary incontinence, chronic urinary incontinence, feeding difficulty, acute confusion, and fever) to indicate whether appropriate steps in care were taken when potentially indicated. Each tracer was reviewed by a panel of clinicians. Because some judgment is involved in whether these steps are always indicated, different values were assigned to the individual steps by a second panel of physicians, who were independent of the project. The summary tracer scores were then used to compare quality changes across time.

Medical record abstraction was performed by trained registered nurses who were not involved in the original study. A week-long session was held to train and test their abstracting skills. In addition, a field manual was developed. The data collection was supervised by staff of the Mountain States Health Corporation.

Analysis

Analyses focused on whether the pattern of pre-post change in nursing homes employing geriatric NPs differed from that in control homes. In some instances the variable of interest was calculated in terms of the change from admission to discharge (or to the three-month time point). The availability of pre-GNP data minimized the effects of any baseline differences between GNP and control homes. The basic analytic tests used analysis of variance for continuous data and chi-square tests for categorical data. The tests for changes in specific medication use employed repeated measures MANOVA. Specific attention was paid to the problem of outliers and their possible effects on mean values.

Newly admitted residents and long-stay residents were analyzed separately. In order to be sure that the analysis addressed the best examples of geriatric nurse practitioner

TABLE 1—Comparisons of Demographic and Clinical Parameters among Study Subjects and Controls

	New Admissions				Long-Stay Patients			
	Pre		Post		Pre		Post	
	GNP	Cont	GNP	Cont	GNP	Cont	GNP	Cont
(N=)	(894)	(981)	(2189)	(2262)	(703)	(606)	(1068)	(1035)
Mean age (yrs)	81.7	82.2	81.7	82.1	83.4	83.7	84.8	84.6
% Admitted from								
Home	22.4	21.1	21.1	21.5	25.6	19.5	21	18.4
Hospital	58.2	61.5	59.3	61.8	50.4	60.5	55.4	60.2
Other	19.4	17.4	19.6	16.7	24.0	20.0	22.5	21.4
% with Diagnosis on Admission								
Dementia	27.5	23.2	22.6	23.2	25.1	24.4	30.4	31.3
Cerebrovascular disease	28.6	25.3	27.0	26.8	31.9	27.8	29.3	26.7
Disease of nervous system	15.3	16.9	20.2	16.2	16.3	15.8	18.5	18.6
Hip fracture	14.0	14.8	15.0	14.4	13.9	15.0	14.7	15.2
Cancer	9.8	10.8	10.5	11.6	6.1	6.6	6.9	7.2
Hypertension	18.4	15.9	21.5	21.1	15.7	17.4	16.7	17.1
Ischemic heart disease	22.4	23.7	22.4	23.2	22.4	24.9	20.6	25.7

effectiveness, the sample was further divided by the extent of GNP role implementation as reflected in interviews conducted with the geriatric nurse practitioners after the evaluation was completed. The GNPs were asked about their role and each practice was classified as full or partially implemented on the basis of the amount of time they reported working on geriatric nurse practitioner tasks. Those working 50 percent or more were considered fully implemented.[7] The results for the fully implemented pairs were calculated in addition to those of the overall sample.

Results

Descriptive Data

As shown in Table 1, data are available on 3,184 cases in the pre-GNP period and on 6,554 cases in the post-GNP period. The pre-GNP ratio of new admissions to long-stay residents is 1.27 for geriatric nurse practitioner cases and 1.63 for control cases. For the post-GNP period the corresponding ratios are 2.0 and 2.2.

Table 1 also compares some basic demographic and historical parameters for the various subgroups. The prevalence of the major chronic disease diagnoses was generally comparable, but among new admissions the GNP residents were more likely to have a diagnosis of dementia in the pre-GNP period and nervous system disease in the post-period. In the post-GNP period, geriatric nurse practitioner residents were more likely to have been admitted from home and less likely to come from the hospital. This same pattern held for long-stayers in both the pre- and post-GNP periods. Not surprisingly, long-stay residents were slightly older than new admissions.

Functional Status

Change in functional status from admission to discharge was calculated as both a continuous score based on the number of domains in which the patient was dependent and a fixed score of change from dependence to independence within each domain. These two approaches allowed for different statistical manipulations but the patterns of overall findings were the same. Table 2 shows that there were few significant differences noted in the extent of change between the geriatric nurse practitioner and control subjects. This table summarizes the differences found among the many analyses performed. Here and in the rest of the tables, the figures presented are the net difference for that variable between GNP and non-GNP from pre to post. The signs indicate the direction of the difference: a positive sign indicates a relative increase in the GNP score compared to the non-GNP from pre to post; a negative sign indicates the reverse.

Case-Mix

Case-mix at the time of admission to the nursing home was measured on the new admissions in three ways: 1) the level of dependence for each of the functional measures shown in Table 2; 2) the frequency of ordered nursing therapies; 3) a case-mix index adapted from that used in Minnesota as a basis for nursing home payment.[8]

Comparisons of individual functional status items on admission showed generally no difference, but in two areas the changes suggested that geriatric nurse practitioner patients were less impaired over time. For the complete sample, more GNP patients were ambulatory and mentally alert. The latter finding was also true for the full implementation subset.

TABLE 2—Change in Functional Status between Admission and Discharge

Functional Status	New Admissions		Long-Stay Patients	
	Complete Sample	Full Implementation	Complete Sample	Full Implementation
Ambulation	−.01	+.01	+.10	+.14
Transferring	+.09*	+.14**	+.08	+.05
Feeding	−.02	−.03	−.02	−.04
Toileting—Urine	+.02	+.02	+.03	+.11
Toileting—Feces	+.01	−.03	+.03	.00
Dressing	+.09**	+.15**	+.04	−.01
Total Number of Dependencies (0–6)	−.07	−.19	−.14	−.14
Mental Alertness	−.02	−.07***	+.05	−.01
Non-Disruptive Behavior	−.03	.00	+.01	+.04

Number = $[(Discharge_{PostGNP} - Admission_{PostGNP}) - (Discharge_{PreGNP} - Admission_{PreGNP})] - [(Discharge_{Post\ Non\text{-}GNP} - Admission_{Post\ Non\text{-}GNP}) - (Discharge_{Pre\ Non\text{-}GNP} - Admission_{Pre\ Non\text{-}GNP})]$

\+ = GNP group relative increase

− = GNP group relative decrease

The possible scores at each point in time ranged from 1 to 5 for ambulation, −6 to +6 for change in total number of dependencies, and 1 and 4 for all other variables.

*.05 < P ≤ .10 **.01 < P ≤ .05 ***P ≤ .01.

These patterns were somewhat different when nursing therapies were used as the basis for case mix. Table 3 shows changes from pre to post for the GNP and control groups. Here the signs indicate an increase or decrease in the ratio. The results reveal a pattern of increased use of several activities that call for patient training (e.g., bowel and gait) and restorative nursing in GNP homes. Moreover, the reductions in use of restraints and even soft diets may reflect an effort by the geriatric nurse practitioners to avoid such approaches to enforced dependency.

Figure 1 compares the patterns of the constructed case-mix index, in which possible scores ranged from 1 to 10. Although both the GNP and non-GNP groups showed pre-

TABLE 3—Change in the Percent of Residents Receiving Nursing Therapies at Admission for New Admits by Level of GNP Implementation

Therapies	Complete Sample	Full Implementation
Decubitus Care	+1.8	+3.9
Foley Catheter	−2.1	−0.1
Bladder Training	+0.2	+3.1
Bowel Training	+3.1*	+1.5
Dressing Change	−1.8	−3.3
Gait Training	+10.8***	+15.1***
I.V. Fluid	+0.3	+0.1
Tube-Feeding	−0.1	−0.6
Ostomy Care	−0.5	+1.5
Restorative Nursing	+15.2***	+23.9***
Oral Suction	+0.3	+0.6
Fracture Care	+2.1	+4.7**
Tracheostomy	+0.1	+0.3
Oxygen	−2.6*	−2.8
Prothesis Care	+1.2*	+0.8
Range of Motion	+0.9	−1.2
Pureed Diet	+0.5	−5.1**
Soft Restraints	−5.3*	−8.8**

Number = $[Post_{GNP} - Pre_{GNP}] - [Post_{Non\text{-}GNP} - Pre_{Non\text{-}GNP}]$

\+ = GNP relative increase from pre to post;

− = GNP relative decrease from pre to post

The possible values for each point in time range from 0–100%, based on a yes-no subject level variable.

*.05 < P ≤ .10 **.01 < P ≤ .05 *** P ≤ .01

post differences, there was little difference in change over time between the groups. The median value for the GNP group was 5.85 in the pre-period and 6.35 in the post-period compared to 5.51 and 5.96, respectively, for the controls. The size of the differences in both means and medians was examined using both ANOVA and various nonparametric tests with the consistent finding of no difference between the GNP and control groups.

Medical Attention

Medical attention was measured in terms of frequency of visits and level of activity. The patterns of change over time are shown in Table 4, where a positive sign reflects a relative increase among GNP groups compared to controls. Among new admissions, there is a substantial reduction in both medication and laboratory orders. The reduction in nursing orders by physicians is reduced when the additive effect of the GNP is considered. By comparing the rates for physicians and the total rates for each service, one can discern a modest pattern of substitution for several services. Interestingly, the rate of physician visits declines for new admissions and increases slightly for the full implementation long-stay residents. There is a consistent pattern of increased physical therapy visits, podiatry, dental care, and occupational therapy, but only the former is significant for the long-stay patients.

Medications

Medication use was examined several ways. The overall number of drugs and doses given was tabulated from the medication sheets, rates were calculated separately for regular administration and PRN use. In addition, dose equivalents were calculated for six common drug classes. These were further analyzed by examining their rate of use with patients who might be expected to require them. The results shown represent changes in the pre-post patterns for differences in use rates from admission to discharge. It is important to appreciate that these summary results can be achieved by either a decrease in post-GNP use or less of an increase.

The summary in Table 5 uses a positive sign to reflect a relative increase for the GNP groups compared to the controls. The numbers shown are differences between GNP and control groups for two-week dosage rates used to summarize the pre-post change in differences from admission to discharge.

The most striking observation is the relative lack of effect. Once again, the differences occur among the long-stay residents of nursing homes rather than newly admitted cases. This overall absence of significant differences resulted from several different phenomena. In some cases, there was truly little shift in usage rates. In others, similar patterns of differences were present for both GNP and controls.

For example, there is much concern about the potential excess use of psychotropic medications among nursing home residents. The net GNP use of psychotropic medication was greater than the controls' for both those with and without disruptive behavior. The GNP mean dose equivalents for psychotropics did not decrease from admission to discharge

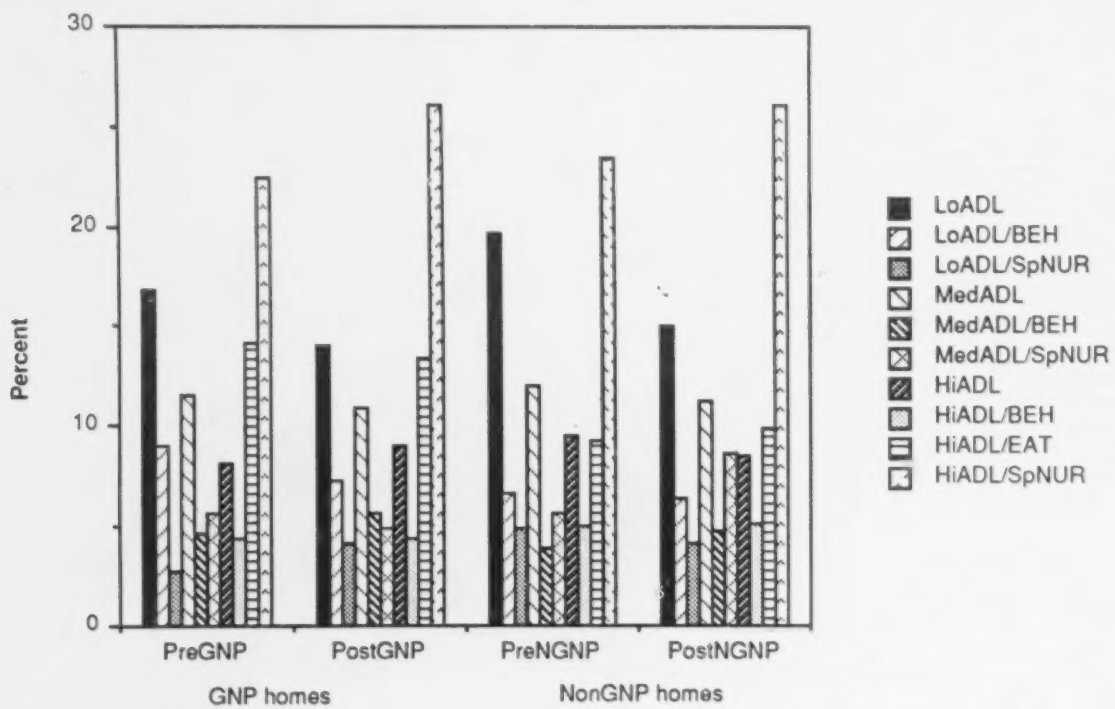

FIGURE 1—Effects of a Geriatric Nurse Practitioner on the Process and Outcomes of Nursing Home Care

ADL = Activities of Daily Living
BEH = Behavior
SpNUR = Special Nursing
EAT = Eating
Lo = Low
Med = Medium
Hi = High
GNP = Geriatric nurser practitioner
NGNP = Control

TABLE 4—Changes in Measures of Medical Attention per 1,000 Patient Days

	New Admissions		Long-Stay Patients		
	Complete Sample	Full Implementation	Complete Sample	Full Implementation	Ranges
Medication Orders					
by Physician	−6.35***	−8.58***	−1.68	−1.57	0–517
by Telephone	−6.52***	−7.39**	−3.56	−.05	0–1049
Total (with GNP)	−9.33***	−10.74**	−3.46	+.52	0–1066
Nursing Orders					
by Physician	−3.61***	−4.45**	−1.16	−1.85	0–289
by Telephone	−1.68	−1.79	−3.99**	−2.94	0–629
Total (with GNP)	−2.44	−2.53	−3.66	−2.95	0–629
Lab Tests and X-Ray					
by Physician	−2.51***	−3.28***	−1.09	−.81	0–414
by Telephone	−2.40***	−2.38	−.72	+1.11	0–333
Total (with GNP)	−3.62***	−3.83*	−.59	+1.11	0–690
Special Orders					
by Physician	+.29	+.94*	−.59	−1.11	0–151
by Telephone	+.07	+1.33**	−.19	−.22	0–200
Physician Visits					
with Examination	−2.21	−3.26	+.60	+1.80	0–1586
without Examination	−.92	−1.04	−1.45	+.72	0–482
Total	−3.22*	−4.44*	−.82	+2.46	0–2069
Podiatry Visits	+.26	+1.25*	+.32	+.09	0–375
Dental Visits	+1.05	+.87	+.03	+.46	0–983
Physical Therapy Sessions	+7.29	+14.38	+14.57**	+17.81**	0–2375
Occupational Therapy Sessions	+.43	+7.42	+2.20	+7.44	0–2475

Number = $[Post_{GNP} - Pre_{GNP}] - [Post_{NON\text{-}GNP} - Pre_{Non\text{-}GNP}]$ Per 1,000 patient days
+ = GNP relative increase from pre to post
− = GNP relative decrease from pre to post
Ranges for any given point in time across all groups are given in the column titled "Ranges".
*.05 < P ≤ .10 ** .01 < P ≤ .05 ***P ≤ .01

for the complete sample, and also ended at a slightly higher level in the post-period than in the pre-period. The control levels did not change from admission to discharge in the post-period and remained at a level below that of the GNP. The GNP overall level of mean dose equivalents for psychotropics in the full implementation sample was lower for the post-period than the pre-period, and decreased within the post-period from admission to discharge. The control level was higher for the post-period than the pre-period.

Tracers

Table 6 shows that in a number of instances the use of GNPs was associated with significant improvements in quality scores. This finding pertains to the newly admitted patients even more than to the long-stay residents. As expected, improvements in quality are somewhat more common among the fully implemented subgroup. In the newly admitted group, positive GNP findings in quality of care are observed in the diabetes and feeding tracers for both, and in the new urinary incontinence tracer in the full implementation sample. Among the long-stay patients, positive effects are found in the congestive heart failure and acute confusion tracers in the full implementation sample and fever in the total sample.

Utilization

There was no difference in the rate of use of emergency rooms for either true emergencies or routine testing. However, the GNP patients were less likely to be admitted to hospital (Table 7). The fully implemented subgroup of new admissions had a significant relative reduction in emergency, elective, and total admissions, based on ANOVA analyses. Total admissions were also significantly reduced for the complete sample of new admissions. The significant relative reduction in emergency and total hospital days occurred among both the full implementation and complete samples of new admissions.

Figure 2 shows the pattern of destinations among those discharged from the nursing homes. Among new admissions, an increased proportion of GNP patients discharged from the nursing home went home. The fully implemented subset of new admissions and long-stay patients were also less likely to be discharged to a hospital.

The discharge patterns were further tested using a series of conditional logistic regression models to control for the effects of other potentially confounding variables including age, source of admission, major diagnoses, payment source, ADL dependencies, and mental status. Separate analyses were performed to examine in sequence the probability of: 1) being discharged dead vs discharged alive or remaining in the nursing home; 2) those alive, going to the hospital or not; 3) those alive, going to the community or not; and 4) those alive and not going to the hospital, going to the community or not. There were no significant differences between GNP and controls during the pre-GNP period. For the post-GNP period the newly admitted GNP residents were less likely to be hospitalized (odds ratio 0.76; confidence interval = 0.59, 0.96) and more likely to be discharged to the community (OR 1.22; CI = 1.02, 1.5). Among the long-stay patients, GNP residents were less likely to be discharged to the hospital. (OR 0.50; CI = 0.34, 0.72).*

Discussion

These data suggest that the geriatric nurse practitioners improved the quality of nursing home care in several impor-

*Data available on request to authors.

TABLE 5—Changes in Medication Use among Nursing Home Patients

	New Admissions		Long-Stay Residents		
	Complete Sample	Full Implementation	Complete Sample	Full Implementation	Ranges
Psychotropic					
All Subjects	+.23	−.19	+3.20***	+3.30*	0–71.4
Disruptive Behavior	+.54	N.A.	+2.53*	N.A.	0–71.4
No Disruptive Behavior	−.21	N.A.	+5.09**	N.A.	0–71.4
Sedatives					
All Subjects	−.14	−.44	+.15	−.29	0–10.0
Disruptive Behavior	−.54	N.A.	+.23	N.A.	0–10.0
Tricylic					
All Subjects	+.19	−.36	+1.14	+4.59*	0–30.0
Disruptive Behavior	+1.53	N.A.	−.01	N.A.	0–20.0
Depressed	+1.11	N.A.	N.A.	N.A.	1.0–20.0
Diuretics					
All Subjects	+.01	−.32	+.08	+.01	0–8.0
Congestive Heart Failure + Hypertension	−.03	N.A.	+1.53	N.A.	0.1–8.0
Lasix					
All Subjects	+.04	+.19	−.11	−.59**	0–7.1
Congestive Heart Failure + Hypertension	+.65	N.A.	−.54	N.A.	0.1–7.1
Digoxin					
All Subjects	−.04	+.02	+.11	+.06	0–3.0
Congestive Heart Failure	−.08	N.A.	+.21	N.A.	0–3.0
Total Doses	−.04	−.26	−.86**	−1.17***	0–43.7
Number of Regular Drugs	+.01	−.01	−.01	−.02	0–1.7
Number of PRN Drugs	−.01	+.00	−.02*	+.03*	0–1.5

Number = $[(\text{Discharge}_{\text{PostGNP}} - \text{Admissions}_{\text{PostGNP}}) - (\text{Discharge}_{\text{PreGNP}} - \text{Admission}_{\text{PreGNP}})] - [(\text{Discharge}_{\text{Post Non-GNP}} - \text{Admission}_{\text{Post Non-GNP}}) - (\text{Discharge}_{\text{Pre Non-GNP}} - \text{Admission}_{\text{Pre Non-GNP}})]$

\+ = GNP group relative increase

− = GNP group relative decrease

Daily dose equivalent ranges for any given point in time across all groups are given in the column titled "Ranges".

*.05 < P ≤ .10 **.01 < P ≤ .05 *** P ≤ .01

TABLE 6—Comparisons of Performance on Tracer Conditions

	New Admissions		Long-Stay Residents		
	Complete Sample	Full Implementation	Complete Sample	Full Implementation	Ranges
Diabetes	+3.00***	+3.59**	+1.25	+3.46	−13.7–
Congestive Heart Failure	−.22	+.36	+1.04	+3.63**	−18.4–
Hypertension	−.27	−.16	+.19	−.09	−10.0–
New Urinary Incontinence	+.86	+1.25**	−.10	−.40	0–6.7
Chronic Urinary Incontinence	+.49	+.91	+.31	+.94	−7.0–
Feeding	+1.00**	+1.89***	+.56	+1.01	0–11.0
Confusion (Acute)	+1.12	−.04	+1.96	+5.11***	0–21.0
Fever (Average)	−.34	−1.20*	+.70	+.89	−2.0–

Number = $[\text{Post}_{\text{GNP}} - \text{Pre}_{\text{GNP}}] - [\text{Post}_{\text{Non-GNP}} - \text{Pre}_{\text{Non-GNP}}]$

\+ = GNP relative increase from pre to post;

− = GNP relative decrease from pre to post

Tracers score ranges for any given point in time across all groups are given in the column titled "Ranges".

*.05 < P ≤ .10 **.01 < P ≤ .05 *** P ≤ .01

tant areas. The benefits are most evident in the tracer studies, which look at the management of specific conditions. The GNP patients show strong indications of closer attention. When these results are compared with the geriatric nurse practitioners' perceptions of where they were having the greatest impact, the patterns validate the GNPs' observations in some areas. For example, the GNPs felt that they had made a substantial impact in drug and laboratory use,[7] and both areas showed reductions among the new admissions. The data suggest a modest substitution of GNP activity for physician activity. The broader measures, such as changes in functional status, similarly do not show much impact although the GNPs thought they were having an effect there.

The geriatric NPs also thought they were permitting an increasing acute case mix. In fact, case mix acuity increased in both the GNP and control homes, but the change is attributable to external forces rather than to adding a geriatric nurse practitioner. Because the timetable for each home was constant within each pair of homes but different across pairs, the effects of changes brought about by shifts in hospital reimbursement and changes in funding rules for nursing home care were minimized in this analysis. The changes in nursing therapies suggest more nursing attention from the GNPs and possibly less use of undesirable approaches to care, such as restraints.

The geriatric nurse practitioner's potential to save total costs of care is suggested by the data on hospital utilization, especially the reduction in hospital days. The savings occur through fewer hospitalizations and less emergency room use. At the same time, GNP patients are more likely to be discharged home from the nursing home.

TABLE 7—Hospital Utilization per 1,000 Patient Days

	New Admissions		Long-Stay Residents		
	Complete Sample	Full Implementation	Complete Sample	Full Implementation	Ranges
Emergency Room Visits					
For Tests	−1.12	−1.81	−.27	−.02	0–453
Not Just for Tests	+.21	−.12	+.14	+.21	0–656
Total	−.09	−1.94	−.14	+.19	0–656
Number of Hospital Admission					
Elective	−.26*	−.54**	+.10	+.08	0–143
Emergency	−.43*	−.96***	−.20	−.16	0–91
Total	−.69**	−1.50***	−.10	−.08	0–143
Hospital Days					
Elective	−.81	−3.21**	+.53	+.33	0–838
Emergency	−5.97**	−8.02**	−1.92	−2.15	0–750
Total	−6.77***	−11.23***	−1.39	−1.82	0–838

Number = $[Post_{GNP} - Pre_{GNP}] - [Post_{Non\text{-}GNP} - Pre_{Non\text{-}GNP}]$
\+ = GNP relative increase from pre to post;
− = GNP relative decrease from pre to post
The range for each point in time across all groups are given in the column titled "Ranges".
* .05 < P ≤ .10 ** .01 < P ≤ .05 *** P ≤ .01

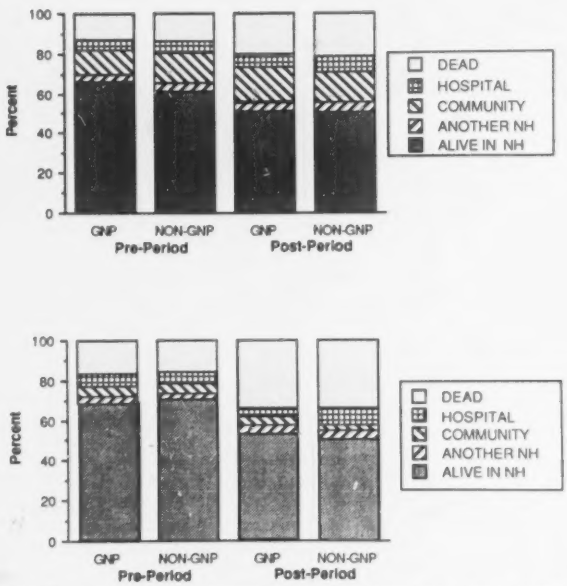

FIGURE 2—Pattern of Discharge Destination by Study Period and Geriatric Nurse Practitioner Status for New Admissions (top figure) and Long-Stay Patients (bottom figure).

The design of the study presents some unavoidable limitations germane to the interpretation of the findings. Although the analysis did control for baseline differences, it cannot control for unmeasured differences between the nursing homes. The homes that undertook this program were adopters and may reflect a more aggressive style of care. The matching did not extend to the level of comparing specific information about staffing ratios.

There were also differences among the GNPs attributable to their individual interests, training, level of physician support, and the administration orientation of their nursing homes. Some of these differences were captured in the interviews with the GNPs, the directors of nursing, and the administrators of the participating homes.[7] To assess the effect of the varied level of implementation of the GNP role, we examined separately those homes where the GNP was primarily performing her intended role and found no consistently different pattern of findings from the total sample.

Taken together, the results of this study suggest that the GNP has a useful and cost-effective role to play in nursing home care. The model tested here relies on the geriatric nurse practitioner as a nursing home employee. It is very possible that GNPs may have an even greater impact if they can operate outside the administrative pressures imposed by this form of sponsorship.

ACKNOWLEDGMENTS

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Major Trauma in Geriatric Patients

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Abstract: Contemporary trauma to the elderly, its severity and associated mortality and morbidity in 111 United States and Canadian trauma centers are described. Three-thousand eight-hundred thirty-three (3,833) trauma patients age 65 years or older are compared to 42,944 injured patients under age 65. Although both groups had equivalent measures of injury severity, the older group had higher case fatality and complication rates and longer hospital stays. The results raise important questions regarding the triage, acute care, accurate prediction of outcome, and hospital reimbursement for the elderly injured patient, with implications for care evaluation, quality assurance, and the long-term viability of trauma centers and systems of care. (*Am J Public Health* 1989; 79:1278–1282.)

Introduction

Today, 11 percent of the United States populace (28.5 million persons) are age 65 or older and by the year 2020 this age group is projected to nearly double, reaching about 51 million.[1] In 1984, unintentional injury accounted for almost 24,500 deaths of persons age 65 and older, a death rate of 86.3 per 100,000 population,[2] more than twice the accidental death rate of all ages (39.3) and for 25 to 44 year olds (35.5).

In 1986, 816,000 person age 65 years and older were discharged from short-stay hospitals with a diagnosis of injury or poisoning.[3] Their average hospital stay was 9.8 days, and the number of bed days exceeded 7.9 million, representing 38 percent of hospital bed days for all patients in which injury was the primary cause of admission.[3] At the 1986 average daily costs in hospital ($500) and in intensive care units (ICU) ($1250–$2000),[4,5] their hospital expenses exceeded $4.4 billion. Increased hospital costs due to longer lengths of stay in hospital and ICU, and more frequent complications,[6,7] are factors contributing to the elderly's disproportionate consumption of approximately one-third of the health care resources expended on trauma care.[8]

Although injury to the elderly consumes a significant proportion of trauma care resources, research has focused on the pediatric and young adult populations. Research has produced quantitative indices of injury severity such as the Trauma Score[9] and the Pediatric Trauma Score,[10] but no comparable scales have been developed specifically for the elderly. A comparative study, however, has shown that the injured elderly differ from their young counterparts in injury etiology, mortality, complications, and cost.[11] A comprehensive picture of injury to the elderly has not emerged and the implications for health care planning remain largely unknown, unquantified, or dimly perceived.

Since 1982, numerous North American hospitals have submitted data and information for seriously injured patients to the Major Trauma Outcome Study (MTOS),[12] sponsored by the American College of Surgeons Committee on Trauma. Data are analyzed periodically and confidential results are sent to participating institutions to support outcome evaluations and quality assurance activities.

This paper compares data from the older (≥65 years) and younger (<65 years) MTOS patients to identify the differences between the two groups and the implications for injury prevention, triage, treatment, outcome evaluation, reimbursement, and additional research.

Methods

Data Base

The MTOS data base includes demographics, cause of injury, injury severity, complications, and outcome information for submitted patients. Detailed text injury descriptions obtained at discharge or death are provided based on surgical examination, x-ray, computerized tomography (CT) scan, and/or autopsy.

Data from 3,833 patients 65 years or older (the "older group") and 42,944 patients less than 65 years (the "younger group") were analyzed. Patient data were submitted from 1982 through 1986 by 111 US and Canadian hospitals participating in MTOS. Ninety-five participants (85.6 percent) reported being Level I or Level II trauma centers.

MTOS participants are required to identify a physician site director and nurse coordinator who are responsible for overseeing data collection, data quality and completeness, and data submission to the study analysis site. The MTOS packet describes the periodic reports prepared for participants, provides operational definitions for each data element, defines the specificity required of anatomic injury descriptions, and suggests sources for data abstraction, i.e., physician and/or operative notes and CT reports.

Collected data are forwarded to a study analysis center in Washington DC, where experienced coders check for data face validity, ensure essential data have been reported, and check the adequacy of submitted injury descriptions for coding. Anatomic injury descriptions are coded according to the International Classification of Disease taxonomy (ICD-9-CM)[13] and for severity using the 1985 version of the Abbreviated Injury Scale (AIS).[14] Data are keypunched, verified, and corrected for analysis. Extensive computer quality control checks are intended to minimize coding errors and to ensure the accuracy, completeness, and consistency of submitted data. When necessary, participants are asked to confirm, clarify, or supplement vague information or missing or unusual data.

Indices of Injury Severity

Patient physiologic status on emergency department arrival is characterized by the Revised Trauma Score (RTS),[15–17] a function of admission values of the Glasgow Coma Scale (GCS), systolic blood pressure (SBP), and

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respiratory rate (RR) (Appendix). The RTS takes values from 0 to 7.8408. Lower RTS values are associated with poorer prognoses.

The RTS replaced the Trauma Score (TS)[9] as the MTOS physiologic severity index. Development of the RTS was motivated by TS limitations; the TS underestimates severity for some head-injured patients and includes variables difficult to obtain in the field (capillary refill and respiratory expansion). The RTS has been shown to be a more reliable predictor of survival/death outcome than the TS.[15]

Anatomic injury severity is measured by the Injury Severity Score (ISS),[18] an index derived from the Abbreviated Injury Scale (AIS)[14] severity scores for individual injuries. The ISS takes values from 1 to 75. Higher scores generally indicate more severe injuries. Studies have shown that the ISS is correlated with mortality; lower correlations have been observed between ISS and length of hospital stay and disability.[18–20]

MTOS survival probability norms are based on the TRISS index.[16,21] TRISS is a severity index which includes the RTS, ISS, and patient age. The mathematical form of the norms is the logistic function $P_s = 1/(1 + e^{-b})$ in which P_s is patient survival probability, $b = b_0 + b_1(RTS) + b_2(ISS) + b_3(AGE)$. Age is 0 for patients <55 years, and 1 for patients ≥55 years of age. The "b"s are regression weights.

TRISS norms are used in the PRE (preliminary outcome evaluation) methodology to support quality assurance activities. Patient ISS and RTS scores are plotted on a scatter diagram (Figure 1). Patients whose coordinates are on the diagonal line (determined by setting b = 0 in the norm equation) are estimated to have a 0.50 survival probability. Coordinates above (below) the line have estimated survival probabilities that are less than (exceed) 0.50. In Figure 1, survivors (L = living) above the line and non-survivors (D = dead) below it represent patients with unexpected outcomes, suggested by the American College of Surgeons to be worthy of peer review.[17]

Statistical Analysis

Statistical comparisons of attributes or outcomes of the older and younger patient samples use the chi-square test for frequencies and z-tests for comparisons of means.

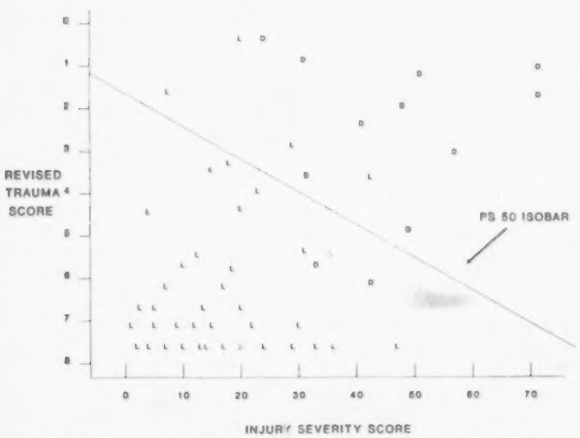

FIGURE 1—Example of a PRE (Preliminary Outcome Evaluation) chart
Key: L = Living
D = Dead

Results

Elderly study patients had a mean age of 76.3 years (sd = 7.8, range 65 to 104 years). Mean age for the younger group was 28.1 years (sd = 13.4, range 1 to 64 years). The male/female ratios for the older and younger groups were .93 and 3.3. The respective mortality rates were 19.0 percent (727/3,832) and 9.8% (4,180/42,860).

Distributions of cause of injury and their associated case fatality rates differed substantially (Table 1). Falls were the most frequent cause of injury to the elderly and had the lowest mortality rate. Motor vehicle accidents, the most prevalent cause of injury in the younger group, was the second most frequent cause of injury to the elderly. Gunshot wounds resulted in the highest case fatality rates in both groups. The elderly showed statistically higher case fatality rates in all etiological categories, except motorcycle injuries and "unknown," for which elderly samples were small.

The severity of anatomic injury and admission physiology as measured by average values of ISS and RTS were not statistically different for the two groups. Mean ISS values were nearly equal for the older and younger groups (13.3 and 13.4, respectively) and mean RTS values were equal (7.0). The difference between the mean ISS values for elderly and young non-survivors (24.1 and 30.2) is substantial, however (6.1, 95% confidence intervals 7.1, 5.1). Elderly non-survivors had a mean RTS at emergency department admission of 4.6, the younger non-survivors had a mean RTS of 2.8 (difference 1.8, 95% CI 2.0, 1.6).

Survival rates for injuries to specific body regions by maximum severity levels are given in Table 2. For example, of patients whose most severe injury was AIS 3 to the thorax, 83.8 percent of the elderly patients survived compared to 95.2 percent of the younger patients. (In Table 2, a patient's outcome is included in the results for each body region sustaining an injury of maximum severity. Thus, the cell entries in a row may not be independent.) The proportion of elderly who died was higher for every severity level/body region combination; differences were especially marked for abdominal injuries (11.5 percent fatality in the younger group and 29.5 percent in the older group). A high proportion of head injuries died in both groups.

Table 3 gives proportions of deaths and complication by ISS interval for the two groups. Mortality increased with higher ISS scores for both patient groups. For every interval, the percentage of deaths of the older group was higher than those of the younger group; the difference was generally greater for ISS values of 25 or more. For ISS values <25

TABLE 1—Mechanism of Injury, Relative Frequency, and Case Fatality Rate

	Ages ≥65		Ages <65	
Mechanism of Injury	% Relative Frequency (N: 3,833)	% Case Fatality Rate	% Relative Frequency (N: 42,944)	% Case Fatality Rate
Fall	40.6	11.7	11.0	6.0
Motor vehicle accident	28.2	20.7	33.5	9.6
Pedestrian hit	10.0	32.6	7.9	13.5
Stab wound	2.6	17.3	11.9	4.7
Gunshot wound	5.5	52.1	13.0	19.5
Motorcycle	0.4	11.8	7.7	11.9
Other	7.0	13.8	14.9	5.4
Unknown	0.3	19.0	0.1	9.8
Total	100	—	100	—

TABLE 2—Percent Survival for Elderly/Younger Patients by AIS Severity and Location of Most Severe Injury

Maximum AIS Severity	Location of Most Severe Body Injury					
	Head and Neck	Face	Thorax	Abdomen, Pelvis	Extremities	External
1	100 /100	94.7/99.7	95.7 /100	100 /100	100 /100	99.7 /100
2	92.6/99.3	95.8/99.6	95.9/99.3	93.3/99.4	94.9/99.5	95.6/99.7
3	66.4/88.1	84.2/98.1	83.8/95.2	77.4/94.7	92.4/96.7	—/98.2
4	54.9/73.2	69.2/76.3	65.6/78.3	65.4/86.7	37.5/88.0	—/77.8
5	26.1/48.0	—/—	23.5/53.8	21.7/66.8	—/69.2	—/37.5

TABLE 3—Percentage of Deaths/Complications According to Injury Severity Score (1985)

ISS Range	Age ≥65	Age <65
0–8	2.9/16.2	0.3/05.8
9–15	6.9/31.1	2.7/17.1
16–24	28.9/48.3	10.5/30.1
25–40	51.4/56.7	29.3/42.9
41–49	73.7/55.3	50.0/53.6
50–74	90.5/61.9	65.2/42.6
75	94.4/38.9	89.9/25.3

(relatively minor or single system injuries), proportions were 9.9 percent and 3.3 percent for older and younger patients, respectively. For patients with an ISS >50 (severe multisystem injuries), most cases in both groups died (92 percent older versus 78 percent younger).

Complications for older (33.4 percent) and for younger patients (19.4 percent) are also very different. Pulmonary complications were most frequent in both groups. Cardiovascular complications were much more common in older patients (9.8 percent) than younger patients (2.1 percent). Complications were substantially more frequent at the less severe ISS intervals for elderly patients than for younger patients.

There were 5.5 percent of blunt-injured MTOS patients and 3.8 percent of patients with penetrating injuries who had "unexpected" outcomes according to TRISS projections (Figure 1). Most striking are those at the extremes of the survival probability scale: 465 unexpected MTOS deaths had estimated survival probabilities exceeding .90. Approximately 30 percent of those patients were ≥65 although the elderly represent only 8.2 percent of all MTOS patients. Most of those patients had no major injury (AIS >3) and had excellent physiology when admitted to the emergency department.

Survivors and non-survivors in the older group had substantially longer average lengths of stay in the hospital and in the ICU than the younger group (Table 4).

Discussion

MTOS findings confirm those of other studies which have shown that the elderly are more vulnerable to injury than their younger counterparts, die more frequently, and have longer hospital stays for injuries of comparable severity.[22,23] This increased vulnerability has in part altered the perception that injury is a disease of the young who, because of the focus on years of productive life lost, have been the beneficiaries of research efforts and health policy thrusts. The magnitude of the problem of injury to the elderly is now being more fully recognized. By 1995, 35 million US citizens will be more than 65 years old. At the current elderly injury mortality rate (86 per 100,000) 30,100 elderly people will die from injuries in 1995. Further, recent studies suggest that trauma is underreported and represents a more prevalent cause of death in the elderly than is officially recognized.[24] The high costs of hospitalizing and treating the elderly are motivating health care providers to examine the efficacy of injury prevention programs and trauma care.

TABLE 4—Mean Length of Hospital and ICU Stay (Days)

Age (years)	Survivors		Non-Survivors	
	Hospital	ICU	Hospital	ICU
≥65	15.7	3.1	9.9	6.0
<65	12.2	2.6	4.2	3.1

There are, however, few reported studies of injury to the elderly and most of those have limited sample sizes. The MTOS thus provides a unique opportunity to study a large number of contemporary, elderly injured. The entire MTOS patient sample represents approximately 0.4 percent of all injured patients hospitalized in the United States, approximately 1.0 percent of all trauma deaths, and 2.0 percent of all in-hospital trauma deaths during the four-year study period. Although the MTOS patient sample is not population based and biased toward the more severely injured patients (more than 85 percent of participating institutions are Level I or Level II trauma centers), the data base focuses on those injuries that are most costly, require a large consumption of health care resources, and perhaps have the greatest impact on the utilization and delivery of trauma care and the formulation of prevention programs.

For example, although falls are the most frequent cause of injury in the elderly, automobile injuries (motor vehicle and pedestrian) result in the most deaths in this age group. The 20.7 percent mortality rate in MTOS for elderly victims of motor vehicle trauma (versus 9.6 percent in the younger group) substantiates the toll of this category of injury. Strategies such as driver improvement programs have been designed to reduce the incidence of motor vehicle deaths in older people, and some state motor vehicle departments and insurance companies have reported a lower incidence of fatalities and collisions from these programs.[25,26] Gunshot wounds, however, are only dimly perceived as a problem in this population. Although the incidence (5.5 percent) is much lower than other injuries, more than half the MTOS gunshot victims age 65 and older died. The rising homicide rate in many cities and higher risk of elderly suicide portend a problem of possibly greater proportions that could require a more focused prevention/treatment strategy.

At the other end of the spectrum are injuries of lesser

severity such as hip fractures, which are often the result (or the cause) of falls in the elderly. The MTOS data base showed an increasing mortality with age in patients whose only injury was a hip fracture. An injury that is easily survivable in the younger population can result in life-threatening consequences for the older group. Identifying such specific causes of vulnerability in the elderly group can target areas for treatment and prevention programs.

A precise characterization of the injury severity of the elderly is as yet unavailable. MTOS outcome prediction results are subject to limitations of the ISS, which have been described,[27] and of the AIS, which assigns severity scores to individual injuries, independent of patient age. Also, investigations of the relation between physiologic scores, such as the RTS, and survival rate among elderly patients have not been reported. Although current injury characterizations include the effect of patient age in a coarse way, factors such as pre-injury illness, disease, or functional level are not considered. These are particularly germane to the elderly group whose susceptibility to complications are well known and who can be at greater risk of dying even when physiology appears normal. The importance of these factors in the characterization of injury to the elderly suggests that more accurate predictions of outcome are possible and necessary. This has been recognized in the pediatric population, for whom a review and reassessment of their AIS scores have already been accomplished.*

MTOS data currently include patient outcomes for survival/death and length of stay in hospital and in ICU. Because of cost and limited resources, most participants cannot perform post-discharge functional assessments, although selected elements of the Functional Independence Measure (FIM)[28] are now being collected. With the limited outcome data, an assessment of the efficacy of trauma center care for the elderly is difficult; refinements are not possible, e.g., determining whether the center benefits elderly injured patients in general or only certain subsets. Such findings could have important system applications such as formulating guidelines for triage. The substantial number of blunt-injured, elderly, "unexpected deaths" identified in MTOS suggests the need to review existing triage criteria endorsed by the American College of Surgeons Committee on Trauma (ACSCOT).[29] Those criteria are based on values of RTS variables and mechanism of injury information. Patient age is mentioned, but unemphatically.

The need for a better understanding of whether the elderly benefit from trauma center care is fueled by the growing emphasis on health care cost containment. Under the prospective payment system, hospital reimbursement is based on DRGs, sets of medically similar diagnoses with approximately equal lengths of stay. Because of the heterogeneity of trauma patients, their hospital stays are bimodal and not well represented by the DRG average. A study of hospital lengths of stay for Maryland patients of all ages with trauma-related DRGs found substantially more variation than for non-injury DRGs considered by many to be the most variable under the system.** Researchers have also noted that among the 469 DRGs, only 96 are adjusted for the elderly. MTOS data suggest the effect of age for all levels of injury severity, and therefore across many diagnoses. Several other analyses suggest that the current system could undermine the financial viability of trauma care systems.[30–32]

MTOS findings indicate that the effect of trauma to the elderly is more serious than to the younger population. The perception of injury as a disease of the young and the use of inadequate physiologic and anatomic indices for this population fail to recognize how trauma affects the elderly and how trauma centers and systems can most effectively be utilized in their care. The elderly, because of their growing numbers, vulnerability to injury, and the significant cost of their trauma care, should be a new focus for trauma research and health policy examinations.

*Susan Baker, PhD, Johns Hopkins University, personal telephone communication.

**Johns Hopkins University proposal to National Center for Health Services Research. Study of Trauma-Related DRGs, 1986. Baltimore, MD: JHU, 1986.

APPENDIX

Revised Trauma Score

Motivated by insights from audits of patients with "unexpected outcomes" identified in the September 1985 MTOS analysis, the Trauma Score (TS) was modified. The Revised Trauma Score (RTS) includes the Glasgow Coma Score (GCS), Systolic Blood Pressure (SBP) and Respiratory Rate (RR).

RTS VARIABLE BREAK POINTS
CODED VALUES

GCS	SPB	RR	Coded Values
13–15	>89	10–29	4
9–12	76–89	>29	3
6–8	50–75	6–9	2
4–5	1–49	1–5	1
3	0	0	0

Each variable takes on coded values from 0 to 4 based upon its raw value. The break-points defining coded values are given below.

Empirically derived weights were obtained for each coded variable from a regression analysis of data from 26,011 consecutive patients from 51 MTOS "norm" institutions. The RTS is defined as

$$RTS = .9368\ (GCS_c) + .7326\ (SBP_c) + .2908\ (RR_c)$$

where the subscript c indicates coded value.

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HRSA Provides Financial Aid to Disadvantaged Students in the Health Professions

More than $5.3 million in grants were awarded recently by the Health Resources and Services Administration (HRSA) to 193 schools in 45 states, the District of Columbia, and Puerto Rico to provide financial assistance to disadvantaged students of medicine, dentistry, or osteopathy. The schools are responsible for selecting students to be assisted. Students from disadvantaged backgrounds who meet exceptional financial need criteria are eligible for support. The aid grants carry no service or financial obligations and may be used for tuition and other reasonable educational expenses. Institutions receiving grants are located in all states except Alaska, Delaware, Idaho, Montana, and Wyoming.

Additionally, grants totaling more than $6.5 million were awarded by HRSA to 298 health professions schools in 46 states, DC, and Puerto Rico to provide scholarships for students in exceptional financial need. Participating schools will award the scholarships to support students of medicine, osteopathy, dentistry, optometry, pharmacy, podiatric medicine, and veterinary medicine during the 1989–90 school year. Those individuals eligible for these scholarships are students in any class year whose resources do not exceed the lesser of $5,000 or one-half the cost of school attendance. The scholarships carry no service or financial obligations. Institutions receiving scholarship funds are located in all states except Alaska, Delaware, Montana, and Wyoming.

For further information about either of these student aid programs, contact Blake Crawford, Health Resources and Services Administration, USPHS, DHHS, 5600 Fishers Lane, Room 14–43, Rockville, MD 20857. Tel: 301/443-3376.

Relation between Vegetarian/Nonvegetarian Diets and Blood Pressure in Black and White Adults

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Abstract: We examined the possible interaction of race and diet on blood pressure (BP) in volunteer Black Seventh Day Adventists compared to volunteer White church members. Height, weight, waist and hip circumference, and resting seated BP were recorded in Black vegetarians (n = 55; age: 54.7 ± 16.9 yrs), Black nonvegetarians (n = 59; 56.1 ± 14.1 yrs), White vegetarians (n = 164; 52.2 ± 16.7 yrs), and White nonvegetarians (n = 100; 52.6 ± 15.6 yrs) attending a regional conference. Forty-four percent of the Black nonvegetarians were medicated hypertensives, compared to only 18 percent of the Black vegetarians, 7 percent of the White vegetarians, and 22 percent of the White nonvegetarians. Black vegetarians exhibited lower age and sex-adjusted systolic BP ($\bar{x}$ = 122.9/74.4 mm Hg) than Black nonvegetarians ($\bar{x}$ = 132.2/75.9 mm Hg). After further adjusting BP for body mass index and waist/hip ratio, the systolic BP among Black vegetarians remained lower (122.8) than Black nonvegetarians (129.7) but higher than that of the Whites who showed no diet-related BP differences. (*Am J Public Health* 1989; 79:1283–1288.)

Introduction

Several dietary constituents including sodium,[1–3] total calories,[4] saturated fat,[5] and alcohol[6,7] have been shown to be positively related to blood pressure (BP). Potassium,[8] calcium,[9,10] magnesium,[11] polyunsaturated fat,[12] and fiber[13] appear to be inversely related to BP. The vegetarian diet tends to be low in the factors positively related to BP, and high in the protective factors (except for calcium in the diet of strict vegetarians who consume no dairy products). Cross-sectional studies of White adults in Australia,[14–16] Israel,[17] and the United States[18–20] have found lower BP among vegetarians than nonvegetarians. Estimates from these studies of the magnitude of the blood pressure differences attributable to diet may be inaccurate because of the lack of random selection of subjects, and group differences in the use of alcohol, caffeine, and tobacco; physical activity; psychosocial stress; weight,[21] body mass index (BMI),[16,19] or anatomic location of body fat.[22] Despite the shortcomings of these observational studies, several experimental studies have been conducted which demonstrate a modest BP-lowering effect when omnivores are switched to a vegetarian diet.[23,24] In these trials a placebo effect cannot be excluded since subjects were not blind to the type of diet consumed.

Whether the vegetarian diet protects against elevated BP has not been studied in US Black adults, a population with a high prevalence of hypertension, possibly due to genetic susceptibility.[25] We compared BP in a group of Black adults following a vegetarian diet to that of a group of Black nonvegetarians. We also explored the possible interaction of dietary preference and race by examining BP in a sample of White vegetarians and nonvegetarians. To reduce the confounding effects of alcohol, caffeine, and tobacco, we studied members of the Seventh Day Adventist Church which encourages avoidance of these substances.

Methods

Subjects

Subjects for this study were volunteers from members of the Seventh Day Adventist Church, who were in attendance at either of two conferences which took place in Michigan during July 1987. The first conference was composed of mostly Black conferees from the midwest region of the US and the second was attended primarily by White church members.

All of the Blacks attending the first conference were requested to participate in this study. Conferees were informed that the purpose of the study was to examine the relation between health habits and health risks. Anonymity and confidentiality were assured; the study was approved by the Committee for the Protection of Human Subjects at Purdue University. Because hypertensives and the elderly might have been more prone to participate, special effort was made to inform the Black conferees that *all* adults were requested to participate, regardless of age or health status. Approximately 50 percent of the Black women volunteered, compared to 20 percent of the Black men. Table 1 provides the demographic characteristics of the 114 Black study participants (89 females, 25 males).

In order to obtain a comparable group of White Adventists, with a sex distribution similar to that of the Black sample, a gender-stratified random sample was recruited from those in attendance at the second conference. From the 400 randomly selected subjects (approximately 80 percent women, 20 percent men) 264 agreed to participate. The characteristics of the 264 White subjects (222 females, 42 males) are also included in Table 1. The average age for all four groups was between 52 and 57 years, and the majority of individuals in each group were long-time church members. The Whites had a longer family history of Adventist church membership than the Blacks, and the White vegetarians had slightly more years of formal education than the other three groups. The average length of adherence to a meatless diet was 20.6 years for Black vegetarians (n = 55) and 27.2 years for White vegetarians (n = 164). Black nonvegetarians (n = 59) averaged 10.1 ± 1.9 ($\bar{x}$ ± SEM) three-ounce servings of meat, poultry, and fish per week, while White nonvegetarians (n = 100) consumed 3.6 ± 0.4 servings of flesh food per week.

Dietary Measures

Subjects were given a comprehensive health questionnaire adapted from previous studies[19,20] and scheduled for measurement of BP and anthropometric characteristics within a four-day period. Subjects categorized themselves as vegetarian (no consumption of meat, fish, or poultry) or nonvegetarian and answered a food frequency questionnaire (Right Byte Nutritionist III, N-Squared Computing, Silverton, Oregon) composed of 141 food items and related food

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TABLE 1—Demographic Characteristics of Black and White Vegetarians (VEG) and Nonvegetarians (NVEG) ($\bar{x}$ ± SEM)

	Black Adventists		White Adventists	
	VEG (n = 55)	NVEG (n = 59)	VEG (n = 164)	NVEG (n = 100)
Age (yrs)	54.7 ± 2.3	56.1 ± 1.8	52.4 ± 1.3	52.6 ± 1.6
Education (yrs)	12.8 ± 0.52	12.8 ± 0.4	13.9 ± 0.2	12.5 ± 0.2
Church member (yrs)	28.7 ± 2.5	29.2 ± 2.6	37.0 ± 1.4	31.2 ± 1.8

groups, to which the subjects reported the amount and frequency of their usual consumption during the previous three months. Included among the listed food items in the instrument are several commercially prepared soy-protein meat analogs frequently used by Seventh Day Adventists. Dietary responses were analyzed using the data base from US Department of Agriculture[26]; macro- and micronutrient intakes were estimated for each subject. Complete dietary data were obtained for 352 of the 378 respondents.

Blood Pressure Measurement

Following a five-minute rest, two seated BP measurements were taken for each subject using a random-zero mercury sphygmomanometer (Hawksley and Sons, Ltd., London) by trained personnel blind to the subject's dietary preference. The second BP measurement was taken two minutes after the first. The first and fifth phase Korotkoff sounds were recorded as systolic and diastolic pressure, respectively. The average of the two BP measurements was used for statistical analyses, regardless of the subject's use or non-use of antihypertensive medications. Subjects were not informed of their BP values until both measurements were recorded.

Anthropometric Measurements

Following the BP measurements, height, weight, triceps skinfold, and waist and hip circumference were assessed. Height without shoes was measured to the nearest one-quarter inch using a stadiometer and converted to centimeters. Subjects were weighed wearing lightweight summer clothing on a balance beam scale. Two triceps skinfold measurements were taken using a Lange skinfold caliper (Cambridge Scientific Industries, Cambridge, MD) and the average was used for statistical analyses. Waist circumference was measured in centimeters using a nonstretchable tape at a level estimated to be midway between the xyphoid process of the sternum and the umbilicus in women, and at the level of the umbilicus in men. Hip circumference was measured in centimeters at the widest hip girth. All circumference measurements were taken without removal of subjects' clothing.

Statistical Analyses

All statistical analyses were made using the Statistical Analysis System (SAS, Cary, NC). Group means of demographic data and anthropometric characteristics were compared using the one-way analysis of variance procedure. Blood pressure data were analyzed with a two-way (diet by race) analysis of covariance, using age, gender, BMI (Wt/Ht^2), waist/hip ratio, and family history of hypertension as covariates. Frequency data were analyzed using the Chi Square statistic. Relationships between variables were examined using simple and multiple correlational analyses.

Results

Intake of macronutrients and several micronutrients of interest are revealed in Table 2. Macronutrient differences between vegetarians and nonvegetarians were as expected, with vegetarians exhibiting lower intakes of protein and fat and higher intakes of carbohydrates. Dietary differences were greater between Black vegetarians and nonvegetarians, than between White vegetarians and nonvegetarians.

As shown in Table 3, Black nonvegetarians exhibited significantly greater body weight, BMI, and waist and hip circumference than the other three groups. The differences

TABLE 2—Estimated Macro and Micronutrient Intake* of Black and White Vegetarian (VEG) and Nonvegetarian (NVEG) Adults ($\bar{x}$ ± SEM)

	Black Adventists		White Adventists	
	VEG (n = 51)	NVEG (n = 49)	VEG (n = 163)	NVEG (n = 89)
Kcalories	2406 ± 130	2504 ± 180	2340 ± 66	2225 ± 92
Protein (g)	66.8 ± 4.9	97.1 ± 9.3	79.9 ± 2.4	80.8 ± 3.4
Protein (kcal%)	12.4 ± 0.4	14.6 ± 0.5	13.0 ± 0.2	14.1 ± 0.3
Carbohydrate (g)	300.5 ± 20.4	321.6 ± 22.5	351.0 ± 9.6	298.3 ± 11.8
Carbohydrate (kcal%)	56.1 ± 1.2	51.0 ± 1.2	57.2 ± 0.6	52.1 ± 7.3
Fat (g)	74.9 ± 5.8	102.7 ± 8.8	82.5 ± 3.2	89.1 ± 4.8
Fat (kcal%)	31.4 ± 1.3	34.5 ± 1.1	29.7 ± 0.6	33.7 ± 7.0
CHOL (mg)	99.4 ± 14.5	294.2 ± 42.0	149.6 ± 9.4	220.1 ± 13.9
Crude Fiber (g)	10.8 ± 1.0	10.2 ± 1.2	13.2 ± 0.6	9.6 ± 0.7
Nondiscretionary Sodium (mg)	2385 ± 197	2833 ± 232	2815 ± 87	2841 ± 133
Potassium (mg)	3721 ± 253	3938 ± 303	4553 ± 129	3945 ± 169
Calcium (mg)	820 ± 79	960 ± 73	1113 ± 45	1103 ± 50
Phosphorus (mg)	1219 ± 101	1534 ± 118	1646 ± 51	1526 ± 59
Magnesium (mg)	368 ± 28	379 ± 29	468 ± 14	364 ± 14

*Nutrient intake estimated from 141-item food frequency questionnaire used to measure habitual intake during the previous three months.

TABLE 3—Anthropometric Characteristics of Black and White Vegetarians (VEG) and Nonvegetarians (NVEG) ($\bar{x} \pm$ SEM)

	Black Adventists		White Adventists	
	VEG (n = 55)	NVEG (n = 59)	VEG (n = 164)	NVEG (n = 100)
Weight (Kg)	70.6 ± 1.6	83.0 ± 2.1	65.2 ± 1.1	71.7 ± 1.6
Body Mass Index (Wt/Ht^2)	26.3 ± 0.5	31.4 ± 0.9	24.6 ± 0.4	27.4 ± 0.6
Waist circumference (cm)	82.4 ± 1.5	92.6 ± 1.8	80.7 ± 1.0	86.8 ± 1.5
Hip circumference (cm)	102.5 ± 1.1	111.8 ± 1.7	101.2 ± 0.8	106.4 ± 1.2
Waist/hip (WHR)	0.80 ± 0.01	0.83 ± 0.01	0.79 ± 0.01	0.81 ± 0.01
Triceps skinfold (mm)	22.1 ± 1.4	32.0 ± 1.9	24.3 ± 0.8	31.1 ± 1.3

between Black vegetarians and nonvegetarians were greater than the differences between the corresponding White groups.

None of the subjects reported current use of cigarettes or tobacco in any form. None of the Black participants consumed alcohol, and only two White nonvegetarians and one White vegetarian reported any use of alcohol. Caffeine consumption averaged less than 45 mg per day for all four groups. Reported levels of occupational physical activity were not different between the four groups. However, Black nonvegetarians reported lower frequency of participation in leisure-time exercise ($\bar{x} \pm$ SEM = 1.4 ± 0.3 sessions per week of aerobic exercise lasting at least 20 minutes per session, compared to the Black vegetarians (2.6 ± 0.3), White vegetarians (2.5 ± 0.2), and White nonvegetarians (2.0 ± 0.2).

Forty-four percent of Black nonvegetarians reported use of physician-prescribed antihypertensive medication compared to only 18 percent of the Black vegetarians, 7 percent of the White vegetarians, and 22 percent of the White nonvegetarians. A higher percentage of Black nonvegetarians (76.3 percent) reported a history of hypertension among their immediate family members (children, siblings, and/or parents) compared to Black vegetarians (55.6 percent), White vegetarians (45.1 percent), and White nonvegetarians (55.0 percent).

Table 4 shows the age- and sex-adjusted systolic and diastolic BP for the four groups. The p-values for main effects of race (Black vs White) and diet (vegetarian vs nonvegetarian) were both less than 0.05 for systolic BP, and there was a weak race by diet interaction ($p = 0.08$). Black vegetarians had lower systolic BP (122.9) than Black nonvegetarians (132.2) but higher than both White vegetarians (113.6) and White nonvegetarians (116.2). Diastolic BP varied with race but not diet. Adding body mass index and waist/hip ratio as covariates, either alone or in combination, attenuated the BP differences between the vegetarians and nonvegetarians, although the patterns are very similar to those adjusted only for age and sex. The addition of family history of hypertension as another covariate in the model did not significantly alter these results for either systolic or diastolic BP.

Table 5 displays the simple correlation coefficients for BP and selected subject characteristics. Age and anthropometric measures were the best single predictors of systolic BP among both the Black and White groups. Waist/hip ratio was a better predictor of both systolic and diastolic BP than was BMI. None of the macro and micronutrient intakes contained in Table 2 (expressed in absolute amounts or per 1000 Kcal) were significantly correlated with BP ($p < 0.01$), nor was the sodium/potassium ratio. For the correlation analyses, the variable "diet" was expanded to four levels by classifying individuals according to four dietary categories based on use of foods derived from animal sources: strict vegetarian; lacto-ovo or lacto vegetarian; prudent diet (use of milk and eggs, and flesh foods no more than three days per week); and omnivorous diet (liberal use of milk and eggs, and flesh foods on a regular basis). Among Blacks, diet was positively

TABLE 4—Adjusted Systolic and Diastolic Blood Pressure for Black and White Vegetarians (VEG) and Nonvegetarians (NVEG) ($\bar{x} \pm$ SEM)

	Black Adventists		White Adventists	
	VEG (n = 55)	NVEG (n = 59)	VEG (n = 164)	NVEG (n = 100)
Blood Pressure Adjusted for Age and Sex				
Systolic BP (mmHg)[a]	122.9 ± 2.3	132.2 ± 2.2	113.6 ± 1.3	116.2 ± 1.7
Diastolic BP (mmHg)[b]	74.4 ± 1.3	75.9 ± 1.2	66.2 ± 0.7	67.7 ± 0.9
Blood Pressure Adjusted for Age, Sex, Body Mass Index, and Waist/Hip Ratio				
Systolic BP (mmHg)[c]	122.8 ± 2.2	129.7 ± 2.2	114.7 ± 1.3	115.0 ± 1.6
Diastolic BP (mmHg)[d]	74.2 ± 1.3	74.8 ± 1.3	66.7 ± 0.7	67.2 ± 0.9

[a]Main effects of diet ($F = 9.8$, $p < 0.01$)
Main effects of race ($F = 43.8$, $p < 0.001$)
Diet × race interaction ($F = 3.03$, $p = 0.08$)
White VEG, White NVEG < Black VEG < Black NVEG (Least Squared Differences, $p < 0.05$)
[b]Main effects of diet ($F = 2.1$, $p = 0.15$)
Main effects of race ($F = 59.4$, $p < 0.001$)
Diet × race interaction ($F = 0$, $p = 0.98$)
[c]Main effects of diet ($F = 3.51$, $p = 0.06$)
Main effects of race ($F = 35.24$, $p < 0.001$)
Diet × race interaction ($F = 3.20$, $p = 0.07$)
White VEG, White NVEG < Black VEG < Black NVEG (Least Squared Differences, $p < 0.05$)
[d]Main effects of diet ($F = 0.28$, $p = 0.59$)
Main effects of race ($F = 49.81$, $p < 0.001$)
Diet × race interaction ($F = 0$, $p = 0.98$)

TABLE 5—Correlations of Systolic Blood Pressure (SBP) and Diastolic Blood Pressure (DBP) with Selected Subject Characteristic of Blacks, Whites, and Combined Races

Variable	Black (n = 114)		White (n = 264)		Combined Races (n = 378)	
	SBP	DBP	SBP	DBP	SBP	DBP
Age	.46[d]	.12	.47[d]	.05	.47[d]	.11[a]
Gender*	.00	−.07	−.18[b]	−.07	−.13[a]	−.10
Body Mass Index	.23[a]	.10	.30[c]	.24[d]	.33[d]	.26[d]
Waist/Hip	.35[d]	.35[d]	.43[d]	.18[b]	.40[d]	.25[d]
Waist	.38[d]	.28[b]	.40[d]	.24[d]	.42[d]	.29[d]
Hip	.24[a]	.08	.23[d]	.20[c]	.27[d]	.21[d]
Triceps skinfold	.08	−.07	.09	.15[a]	.09	.07
Exercise frequency**	−.18[a]	−.09	.05	−.02	−.06	−.07
Diet†	.23[a]	.10	.07	.09	.18[c]	.15[c]
Family history of HT††	.01	.00	.16[b]	.18[b]	.16[b]	.17[c]
Race‡	—	—	—	—	−.34[d]	−.39[d]
Education	−.18[a]	−.18[a]	−.20[b]	.05	−.21[d]	−.09
Calcium	.13	.17	−.08	.00	−.06	−.02
Sodium	.07	.13	−.05	.02	−.03	.03
Sodium/Potassium	−.01	.10	−.01	−.05	.00	.01
Polyunsaturated fat/ saturated fat	−.12	−.13	.06	.09	.04	.06

[a] $p < 0.05$
[b] $p < 0.01$
[c] $p < 0.001$
[d] $p < 0.0001$
*Gender (male = 1, female = 2)
**Exercise Frequency (no exercise = 0, times per week = 1, 2, 3 . . .)
†Diet (strict vegetarian = 1, lacto-ovo vegetarian = 2, prudent diet = 3, omnivorous diet = 4)
††Family history of hypertension (no = 1, yes = 2)
‡Race (Black = 1, White = 2)

associated with BP in the simple correlation analyses, but was not among the Whites, possibly owing to their more homogenous diet.

Stepwise multiple regression analysis was performed using variables significantly related to BP in the simple correlation procedure as well as dietary intake of fat, total fat, sodium, potassium, sodium/potassium ratio, calcium and magnesium. Age, race, waist circumference, and family history of hypertension were the best predictors of systolic BP for all subjects. Race, waist circumference, and family history of hypertension were the best predictors of diastolic BP. The unadjusted positive correlation between diet and BP was partly explained by the effects of waist circumference and waist/hip ratio. None of the specific nutrients significantly contributed to the prediction of blood pressure in the multiple regression analysis.

Discussion

The results obtained from the Black participants in this study parallel those of several previous studies of White adults showing that vegetarians exhibit lower BP than nonvegetarians independent of differences in anthropometric characteristics.[14–17] Systolic BP differences between the two Black groups would likely have been even greater had not significantly more Black nonvegetarians had their high BP attenuated by use of antihypertensive medications compared to the Black vegetarians. The lack of significant differences in BP between the White dietary groups may have been the result of inadequate dietary heterogeneity. White Adventist nonvegetarians eat a more prudent diet than the usual omnivore, as evidenced in this study by their average consumption of only three to four, three-ounce servings of meat, fish, or poultry per week, fewer than 300 mg dietary cholesterol per day, and less than 34 percent of calories as fat. Indeed their diet is more similar to the vegetarian diet than that of omnivores in previous studies.

Dietary differences were greater between the Black vegetarians and nonvegetarians which may explain the lower systolic blood pressure in the former, even after adjustment for anthropometric characteristics. The positive association between systolic blood pressure and diet (at four levels of animal food consumption) provides supporting evidence. Possibly the systolic blood pressure difference between Black dietary groups was related to better blood pressure control among the medicated Black hypertensive vegetarians ($\bar{x}$ BP = 136/74) compared to Black nonvegetarians on antihypertensive therapy ($\bar{x}$ BP = 145/79). However, all but two hypertensive subjects (one vegetarian and one nonvegetarian) reported adherence to prescribed medications. Other studies suggest the vegetarian diet may reduce the effective drug dosage for medicated hypertensives.[27]

Despite similarities in age, anthropometric characteristics, exercise habits, and reported levels of stress, Black vegetarians still exhibited a higher prevalence of diagnosed hypertension and a higher average systolic and diastolic blood pressure than White vegetarians. In comparison to White vegetarians, the Black vegetarians had significantly lower mean intakes of potassium, calcium, vitamin D, and magnesium, all of which have been reported to be inversely related to blood pressure.[8–11,28,29] However, the correlational analyses provide no support for their contribution to the blood pressure differences. The absence of a clear nutrient effect, while somewhat puzzling, has also characterized other studies of vegetarian diet and blood pressure. Limitations of questionnaire data are an obvious source of some misclassification, but even controlled dietary trials by Sacks, *et al*,[30] have had difficulty in identifying specific foods and nutrients responsible for the lower blood pressure among vegetarians. They concluded that lower intakes of total fat,

saturated fat, and animal protein, and higher intakes of carbohydrates, polyunsaturated fatty acids, and vegetable protein do not alone explain the lower blood pressure in vegetarians compared to nonvegetarians. In addition, based on experimental findings, Margetts, *et al*,[31] have concluded that the vegetarian diet does not affect blood pressure by way of its high fiber content.

The Australian research group, which has performed numerous studies regarding diet and blood pressure, finds strong evidence for an independent blood pressure-lowering effect of a vegetarian diet, but considerable uncertainty as to the nutrients responsible for this effect.[32] Multiple nutrients may act in concert, with the effect on blood pressure due to the totality of the vegetarian diet, rather than single dietary constituents. Cross-sectional analyses commonly find little or no correlation between specific nutrients such as sodium and blood pressure.[33] This may be partly explained in our study by attempting to relate only current nutrient intake (previous three months) to blood pressure when habitual nutrient intake over several years or even decades would appear to have a stronger influence on blood pressure. The inclusion in the regression analysis of BP values achieved by use of antihypertensive medication could also mask significant dietary correlates of blood pressure. Although attempts were made to minimize non-dietary group differences by sampling only from among Adventist church members, possibly self-selected vegetarians differ systematically from nonvegetarians in attributes not measured in this study, which might be important determinants of blood pressure.

Among both Black and White Adventists, central adiposity as estimated by waist circumference or waist/hip ratio was a stronger predictor of blood pressure than was BMI or peripheral adiposity estimated by triceps skinfold or hip circumference. Dietary preference was related to fat patterning with vegetarians of both races exhibiting less central deposition of fat as estimated by waist circumference and waist/hip ratio. The lower weight and adiposity among Black vegetarians compared to nonvegetarians appears to be at least partly explained by differences in caloric intake and frequency of exercise. However, White vegetarians were significantly leaner than White nonvegetarians despite similar intake of calories and reported exercise frequency. Factors other than total caloric intake and physical activity are known to influence body weight including dietary composition, resting metabolic rate, postprandial thermogenesis, and adaptive thermogenesis.[34] Virtually all studies to date have reported vegetarians to weigh significantly less than nonvegetarians, yet the specific reasons for this difference are unclear. A recent study from our laboratory suggests the lower body weight of vegetarians is not due to higher resting metabolic rate or postprandial thermogenesis compared to nonvegetarians.[35]

In summary, it appears that while there is a relationship between dietary preference and blood pressure in Black adults, the vegetarian diet does not completely attenuate their apparently greater susceptibility to hypertension compared to Whites. However, from a public health perspective, the vegetarian diet appears to have implications for the primary prevention of hypertension and obesity for both Black and White adults. Because such a diet is not easily adopted by the majority of adults in the US, future research must continue the search for the specific dietary factors which afford protection against abnormal elevation of both blood pressure and body weight.

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Births to Older US Women: NCHS Report

Detailed statistics on first births to older US women are now available from the National Center for Health Statistics in a new report, *Trends and Variations in First Births to Older Women, 1970–86*. Statistics in the report focus on the decline in childbearing among well-educated women in their twenties, and the extent to which these women have delayed motherhood. Maternal and infant health characteristics are also analyzed. Highlights of the report include:

- First births to women in their thirties have increased substantially since 1970, quadrupling their levels in a 16-year period. Throughout the 1970s, the increases were largest for women ages 30–34 years. Since 1980, the impact of this phenomenon has shifted to older women. Between 1980 and 1986, the increases in first births and first-birth rates were much greater for women ages 35–39 years.
- During this period of rapidly rising first-birth rates for women in their thirties, rates for women in their teens and early twenties declined.
- The number of women in their thirties grew substantially during this period—all born in the post World War II baby boom years.
- Women aged 30 and older accounted for 15 percent of all first births in 1986 compared with just 4 percent in 1970.
- Unmarried women accounted for about one in 10 first births to mothers ages 30–34 in 1986 compared with one in 20 a decade earlier.
- The educational attainment of women having their first child at ages 30 and older has increased substantially, and is higher than for women in any other age group. During the 1980s, close to one-half of the older first-time mothers have been college graduates, compared with one-fourth of comparable women in 1970.
- Greater educational attainment is associated with more timely receipt of prenatal care and a better outcome in terms of birthweight. More than 90 percent of mothers in their thirties who were college graduates began prenatal care in the first trimester. Only 5.8 percent of first births to college graduates ages 30–34 were low birthweight in 1986 compared with 7.5 percent of all first births to mothers in this age group.

For further information or copies of the report, contact: Sandra Smith, Public Information Officer, NCHS, 301/436-7135; or Stephanie Ventura, Statistician, Natality, Marriage and Divorce Statistics Branch, 301/436-8954.

Sample Size and Power Based on the Population Attributable Fraction

WARREN S. BROWNER, MD, MPH, AND THOMAS B. NEWMAN, MD, MPH

Abstract: Most methods for calculating sample size use the relative risk (RR) to indicate the strength of the association between exposure and disease. For measuring the public health importance of a possible association, the population attributable fraction (PAF)—the proportion of disease incidence in a population that is attributable to an exposure—is more appropriate. We determined sample size and power for detecting a specified PAF in both cohort and case-control studies and compared the results with those obtained using conventional estimates based on the relative risk. When an exposure is rare, a study that has little power to detect a small RR often has adequate power to detect a small PAF. On the other hand, for common exposures, even a relatively large study may have inadequate power to detect a small PAF. These comparisons emphasize the importance of selecting the most pertinent measure of association, either relative risk or population attributable fraction, when calculating power and sample size. (*Am J Public Health* 1989; 79:1289–1294.)

Introduction

In most observational studies the relative risk (RR)—the incidence rate of a disease in persons exposed to a risk factor divided by the incidence rate in persons without the exposure—measures the association between exposure and disease. The relative risk is often used to indicate the strength of that association; in general, the stronger an association, the more likely that it is causal. The goal of most observational studies is the discovery of causal relationships, and sample size and power are generally calculated to detect a given relative risk.

In other circumstances, however, an investigator may be more interested in determining the overall impact of an association than its causal strength. For example, an investigator studying congenital heart disease might want to have sufficient sample size to detect those exposures associated with 5 percent or more of the cases. For this purpose, the population attributable fraction (PAF)—the proportion of disease incidence in a population that can be attributed to a particular exposure—is a more useful measure than the relative risk. Assuming that the exposure is a cause of the disease, the population attributable fraction indicates the potential public health importance of an association. The higher the PAF, the greater the potential for reducing the incidence of the disease by modifying that exposure.[1]

The population attributable fraction[2,3] (also called attributable risk,[4] population attributable risk proportion,[5] etiologic fraction[6]) is defined as the excess incidence of a disease that is due to an exposure in a population, divided by the total incidence of the disease in the population. If I is the incidence rate of a disease in the population, and I_{un} is the incidence rate in those without the exposure, then the excess incidence of the disease due to the exposure is $I - I_{un}$, and:

$$PAF = \frac{(I - I_{un})}{I} \qquad (1)$$

For example, if the annual incidence of lung cancer in middle-aged men in the United States is 46 per 100,000, and the incidence in non-smokers is 10 per 100,000, then 78 percent of the lung cancer in this age group can be attributed to current smoking. In its more familiar form, as it was introduced 35 years ago by Levin,[7] the PAF equals:

$$\frac{p_x\,(RR - 1)}{p_x\,(RR - 1) + 1} \qquad (2),$$

where p_x is the prevalence of the exposure (risk factor). Since the RR of lung cancer in smokers is about 10, and the prevalence of smoking among middle-aged men is about 0.40,

$$PAF = \frac{.4\,(10 - 1)}{.4\,(10 - 1) + 1} = 0.78.$$

The PAF depends upon both the exposure prevalence and the relative risk: at any given population attributable fraction, rare exposures must be associated with large relative risks to have even a modest PAF. (The relationship between exposure prevalence, relative risk, and PAF is discussed further in several of the references.)[8–10] At an exposure prevalence of 5 percent, for example, the relative risk must be 6 in order to have a PAF of 20 percent. Conversely, a common exposure can be responsible for the same PAF with a small relative risk: at an exposure prevalence of 50 percent, a relative risk of only 1.5 has a PAF of 20 percent.

As we will show, at the same PAF, it turns out to be "easier" (power is greater, or sample size is smaller) for a study to detect a rare exposure with a large relative risk than to detect a more common exposure with a smaller relative risk. For example, in a case-control study with one control per case, at one-sided alpha = 0.025 and power = 0.80, at a PAF of 20 percent, 64 subjects are required when the exposure prevalence is 5 percent (and RR = 6), whereas 408 subjects are needed when the exposure prevalence is 50 percent (and RR = 1.5).

We have estimated sample size and power for observational studies based on the detection of a given population attributable fraction for a binary (yes/no) exposure, and compared them with conventional estimates based on detecting a given relative risk.[11–13] These methods complement currently available methods for determining the confidence interval for a population attributable fraction once a study has been completed.[14,15]

Methods

To estimate the sample size required to detect a specified population attributable fraction, we developed sample size

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formulas, using the PAF, rather than the relative risk or odds ratio, as the measure of association. The null hypothesis (H_0) is that the PAF = 0; the alternative hypothesis (H_1) is that the PAF ≥ 0. We applied the Fleiss-Tytun-Ury approximation for the continuity correction.[11]

Sample Size for Cohort Studies

Using the risk ratio (RR) as the measure of association, the number of unexposed subjects, m′, required in a cohort study, prior to applying the Fleiss-Tytun-Ury approximation for the continuity correction, is:[12]

$$m' = \frac{[Z_\alpha \sqrt{\bar{I}(1-\bar{I})} + Z_\beta \sqrt{p_x I_{un}(1-I_{un}) + (1-p_x) RR I_{un}(1-RR I_{un})}]^2}{p_x [(RR-1) I_{un}]^2}$$

(We assume that p_x, RR, I, and I_{un} in the cohort estimate their respective values in the population.)

We can replace RR with PAF by solving equation 2:

$$RR = \frac{PAF}{p_x (1 - PAF)} + 1 \qquad (3),$$

and can replace I_{un} by solving equation 1:

$$I_{un} = I (1 - PAF).$$

Note that:

$$RR - 1 = \frac{PAF}{p_x (1 - PAF)},$$

and

$$(RR-1)\, I_{un} = (1/p_x)\, I\, PAF.$$

Thus:

$$m' = \frac{p_x (Z_\alpha \sqrt{I(1-I)} + Z_\beta \sqrt{p_x T + (1-p_x) U})^2}{PAF^2 I^2}$$

where

$$T = [I (1 - PAF)] (1 - [I(1 - PAF)]),$$

and

$$U = p_x^{-2}(I[PAF + p_x(1-PAF)])[p_x - (I[PAF + p_x (1-PAF)])].$$

Applying the approximate continuity correction, the corrected number of unexposed subjects required, m, is obtained:

$$m = m' + \frac{1}{p_x (RR - 1) I_{un}},$$

$$= m' + \frac{1}{PAF\, I}.$$

Thus the total number of subjects, N, required in the cohort would be:

$$N = \frac{m}{(1 - p_x)}.$$

Sample Size for Case-Control Studies

We assume that the disease is rare, or that incidence-density sampling of controls has been used,[16] and that cases and controls represent the population with and without the disease, respectively. Then p_x, the proportion of controls with a binary exposure, is approximately equal to the proportion of the population with the exposure; p_c, the proportion of cases with an exposure, is approximately equal to the proportion of incident cases in the population with the exposure; and the odds ratio (OR) is approximately equal to the rate ratio for the disease. Let P equal the overall proportion of subjects (cases and controls) exposed, and b represent the control:case ratio.

The total number of cases, c′, required in a case-control study, prior to applying the Fleiss-Tytun-Ury approximation for the continuity correction, is:[13]

$$c' = \frac{(Z_\alpha \sqrt{(1+1/b)P(1-P)} + Z_\beta \sqrt{(p_c)(1-p_c) + p_x(1-p_x)/b})^2}{(p_c - p_x)^2}.$$

When the odds ratio (OR) serves as the measure of association:

$$p_c = \frac{p_x\, OR}{1 + p_x (OR - 1)}.$$

By analogy to equation 3:

$$OR = \frac{PAF}{p_x (1-PAF)} + 1.$$

Since by definition:

$$OR = \frac{p_c (1 - p_x)}{p_x (1 - p_c)},$$

thus

$$p_c = PAF + p_x (1 - PAF),$$

when the PAF serves as the measure of association.

With b controls per case:

$$P = \frac{p_c + b\, p_x}{b + 1}$$

$$= \frac{PAF\ (1 - p_x) + p_x\ (b + 1)}{(b + 1)}.$$

Thus:

$$c' = \frac{(Z_\alpha \sqrt{X} + Z_\beta \sqrt{Y})^2}{PAF^2 (1 - p_x)^2},$$

where

$$X = \frac{[PAF(1-p_x) + p_x(b+1)]\ (b+1-[PAF\ (1-p_x) + p_x\ (b+1)])}{b\ (b+1)},$$

and

$$Y = [PAF + p_x\ (1-PAF)]\ (1-p_x)\ (1-PAF) + [p_x\ (1-\ p_x)/b].$$

Applying the Fleiss-Tytun-Ury correction, the corrected number of cases, c, would be:

$$c = c' + \frac{b + 1}{b \mid p_c - p_x \mid}$$

$$= c' + \frac{b + 1}{b\ PAF\ (1 - p_x)}$$

Results

We compared the results of sample size and power calculations based on a specified population attributable fraction with the usual ones based on the relative risk or odds ratio for cohort and case-control studies. We used a one-tailed alpha of 0.025 (Z = 1.96).

Table 1 presents the number of subjects required in a cohort study to detect a given population attributable fraction for different combinations of the incidence of a disease and the prevalence of an exposure. If the incidence of a disease is 5 percent, for example, 59,371 subjects would be necessary to detect an exposure of prevalence 0.80 that was responsible for 10 percent of the disease (the corresponding relative risk is 1.14).

Table 2 presents the number of cases required in a case-control study to detect a given PAF for different combinations of the number of controls per case (the control: case ratio) and the prevalence of an exposure among the controls. If there are 2 controls per case, for example, 28 cases would be needed to detect an exposure of prevalence 0.30 that was responsible for 50% of the disease (an odds ratio of 4.33).

Comparison with Power and Sample Size Based on Relative Risk or Odds Ratio

Figure 1 shows the familiar relationship between power and exposure prevalence (P_x, the proportion exposed to a risk factor) at several relative risks for a reasonably large cohort study (chosen to be the same size (N = 2179) and disease incidence (I = 0.08) as the Framingham, Massachusetts study of coronary heart disease in middle-aged men[17]). Note that as the prevalence of the exposure approaches approximately 0.50, power increases. Thus the more common an exposure, up until about half of the cohort has been exposed, the greater the power to detect a given relative risk. For all but the rarest (or most common) exposures, there would be enough subjects in the cohort to have adequate power to detect relative risks of 1.8 or more.

By contrast, in Figure 2 power is plotted as a function of the exposure prevalence for several values of the population attributable fraction. Power is maximal as prevalence approaches zero (and the corresponding relative risk approaches infinity), and decreases rapidly as prevalence increases (and the corresponding relative risk gets smaller). Thus even a study with more than 150 events among more than 2000 subjects would have only limited power to detect population attributable fractions of 15 percent or less for common exposures ($P_x > 0.40$). On the other hand, the study would have sufficient power to detect uncommon exposures ($P_x < 0.10$) associated with population attributable fractions as small as 5 percent.

Figure 3 shows the familiar relationship between: 1) the number of cases required in a case-control study; 2) Px, the proportion of controls exposed to a risk factor; and 3) the odds ratio (with two controls per case, at a power of 0.80). The number of cases required to detect a given OR is minimized as P_x approaches about 0.50. For example, a

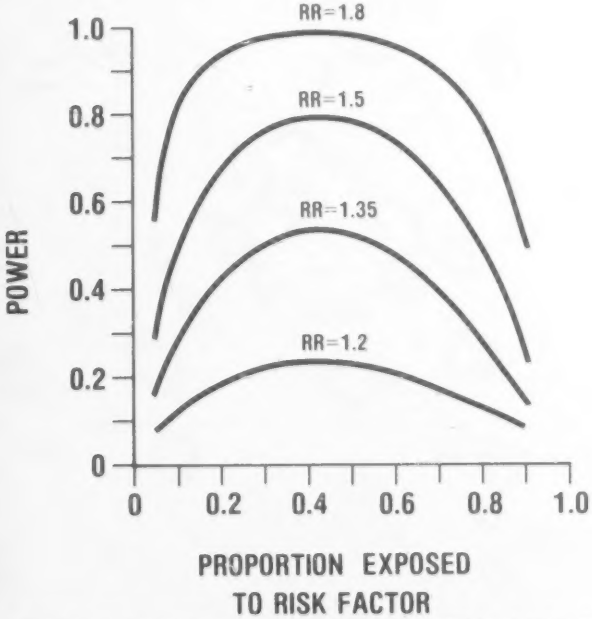

FIGURE 1—Power to detect a given relative risk (RR) for CHD as a function of the proportion of cohort exposed to a risk factor (P_x) in a cohort study (with N = 2179, incidence proportion = 0.08, and one-tailed alpha = 0.025).

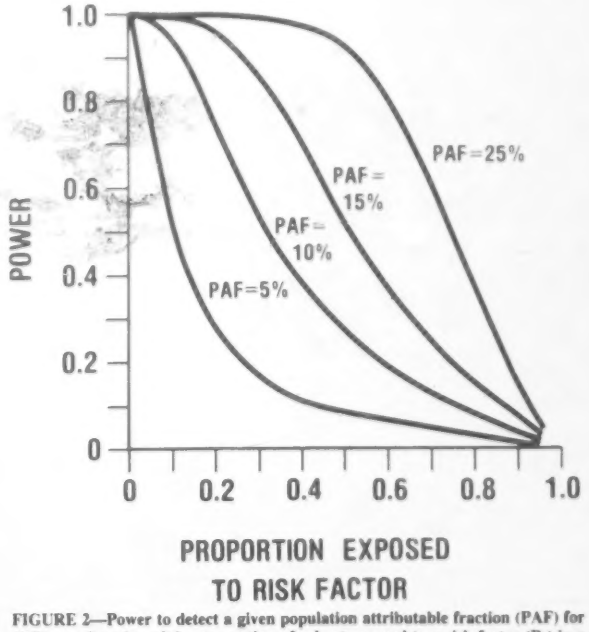

FIGURE 2—Power to detect a given population attributable fraction (PAF) for CHD as a function of the proportion of cohort exposed to a risk factor (P_x) in a cohort study (with N = 2179, incidence proportion = 0.08, and one-tailed alpha = 0.025).

TABLE 1—Total Sample Size Required in Cohort Study, Based on Population Attributable Fraction (PAF), Prevalence of Exposure (P_x), and Incidence of Disease (at power = 0.80, one-sided alpha = 0.025).*

P_x	.05	.10	.15	.20	.25	.30	.40	.50	.60	.70	.80	.90
PAF												
	22332	40622	60879	83604	109330	138713	212138	314906	469040	725917	1239661	2780867
5%	4246	7770	11665	16033	20975	26620	40726	60467	90076	139420	238105	534158
	1984	3654	5514	7586	9931	12609	19299	28662	42705	66107	112911	253319
	6920	11579	16661	22331	28730	36029	54247	79725	117925	181578	308870	690728
10%	1299	2205	3186	4277	5509	6913	10416	15315	22660	34898	59371	132787
	595	1033	1501	2021	2606	3273	4938	7264	10752	16563	28184	63044
	3631	5750	8025	10545	13380	16607	24646	35877	52705	80737	136787	304918
15%	671	1089	1530	2017	2564	3185	4733	6895	10133	15527	26313	58665
	300	506	718	951	1212	1508	2244	3272	4812	7376	12503	27883
	2340	3561	4853	6274	7865	9672	14165	20430	29810	45430	76654	170304
20%	426	670	923	1198	1506	1854	2720	3927	5734	8742	14756	32791
	185	309	431	563	711	877	1290	1865	2725	4157	7018	15601
	1681	2482	3317	4230	5248	6400	9257	13234	19183	29082	48863	108187
25%	300	464	629	806	1004	1226	1778	2545	3692	5600	9412	20846
	126	212	292	378	473	580	843	1209	1755	2664	4481	9928
	1291	1861	2448	3084	3791	4589	6562	9301	13393	20198	33790	74544
30%	226	346	462	587	724	879	1260	1789	2579	3891	6513	14373
	91	156	214	274	341	415	597	850	1227	1853	3103	6852
	1036	1466	1902	2372	2892	3477	4918	6915	9892	14839	24715	54317
35%	178	271	358	450	552	665	944	1330	1905	2860	4766	10480
	68	121	165	210	259	314	448	632	907	1363	2273	5000
	859	1196	1535	1896	2295	2742	3840	5356	7613	11359	18833	41226
40%	145	219	288	359	437	524	737	1030	1467	2190	3634	7959
	52	96	132	167	205	247	349	490	698	1044	1734	3801
	730	1003	1273	1561	1876	2229	3092	4281	6047	8973	14806	32274
45%	120	182	238	295	357	426	593	824	1165	1731	2858	6234
	39	79	108	137	167	200	281	392	555	826	1365	2979
	632	858	1080	1315	1571	1856	2552	3507	4923	7265	11928	25887
50%	102	155	201	248	298	354	490	675	949	1402	2304	5003
	—	66	91	114	139	166	232	321	452	669	1101	2393

*Use upper number if overall incidence of disease is 0.01 (1 per 100); middle number if incidence is 0.05; and lower number if incidence is 0.10.

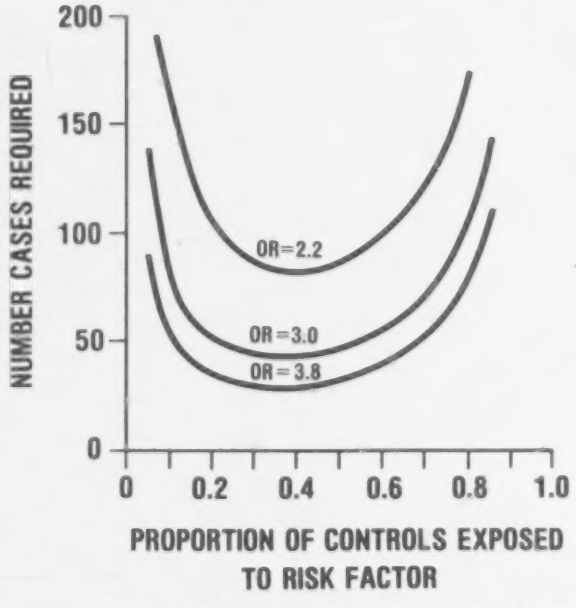

FIGURE 3—Number of cases required to detect a given odds ratio (OR) as a function of the proportion of exposed controls (P_x), at control:case ratio = 2:1, power = 0.80, one-tailed alpha = 0.025.

case-control study with 91 cases and 182 controls—the same size as a study of risk factors for hip fracture reported by Paganini-Hill, *et al*,[18]—would have a power of 0.80 to detect an odds ratio of about 2.2 for reasonably common exposures ($0.30 < P_x < 0.55$). As expected, however, if only 1 in 20 controls was exposed, the study would have adequate power to detect only odds ratios of 3.8 or more.

When plotted as function of P_x, for any given value of the population attributable fraction (Figure 4), the number of required cases is minimized as P_x approaches zero. For a rare exposure like diabetes ($P_x = 0.05$[18]), a study with about 90 cases and 180 controls, for example, would have a power of 0.80 to detect a population attributable fraction of 15 percent (a corresponding relative risk of 4.5). But for a more common exposure like use of sedative-hypnotic medications ($P_x = 0.52$[18]), the study would have adequate power to detect only population attributable fractions of more than 40 percent (relative risks of 2.3 or more).

Discussion

The relationship between power, relative risk and population attributable fraction may seem paradoxical. Even though the power to detect a given relative risk (or odds ratio) increases as an exposure becomes more common (up to about 50 percent), the power to detect a given population attributable fraction actually decreases. This occurs because the population attributable fraction depends upon both the prev-

TABLE 2—Number of Cases Required in Case-Control Study, Based on Population Attributable Fraction (PAF), Prevalence of Exposure (P_x) among Controls, and Number of Controls per Case (at power = 0.80, one-sided alpha = 0.025).*

P_x	.05	.10	.15	.20	.25	.30	.40	.50	.60	.70	.80	.90
PAF												
	517	877	1280	1733	2247	2834	4302	6356	9438	14575	24849	55668
5%	373	644	948	1289	1675	2117	3222	4768	7087	10953	18684	41876
	285	504	748	1022	1332	1687	2573	3815	5676	8779	14984	33600
	173	262	361	472	598	742	1102	1606	2362	3621	6141	13699
10%	123	190	265	349	444	553	824	1205	1776	2727	4630	10337
	91	146	206	274	351	439	658	964	1424	2190	3722	8319
	96	134	177	226	281	343	500	720	1049	1598	2697	5991
15%	67	97	129	166	208	255	374	541	790	1206	2039	4535
	49	73	100	130	163	202	298	433	634	971	1643	3660
	64	85	109	135	166	200	287	408	590	892	1498	3314
20%	45	61	79	99	122	149	215	307	445	675	1136	2517
	32	46	61	77	96	117	171	245	358	544	918	2038
	47	60	75	92	111	132	187	263	377	566	946	2085
25%	33	43	54	67	82	98	140	198	285	430	719	1589
	23	32	41	52	64	77	111	158	229	347	583	1290
	36	45	56	67	80	95	132	183	261	390	648	1423
30%	25	32	40	49	59	70	98	138	197	296	494	1088
	18	24	30	38	46	55	78	111	159	240	401	886
	29	36	43	51	61	71	98	135	191	284	469	1027
35%	21	26	31	38	45	53	73	102	145	216	359	788
	15	19	24	29	35	41	58	82	117	175	292	643
	24	29	35	41	48	56	76	104	145	215	354	772
40%	17	21	25	30	35	41	57	78	111	164	272	594
	12	15	19	23	27	32	45	63	89	133	222	487
	21	24	29	33	39	45	60	82	114	168	275	598
45%	15	17	21	24	29	33	45	62	87	129	212	462
	11	13	16	19	22	26	36	50	70	105	173	380
	18	21	24	28	32	37	49	66	92	134	220	475
50%	13	15	17	20	24	28	37	50	70	103	170	369
	9	11	13	16	18	21	29	40	57	84	139	303

*Use upper number if there is 1 control per case; middle number if there are 2 controls per case; and lower number if there are 5 controls per case.

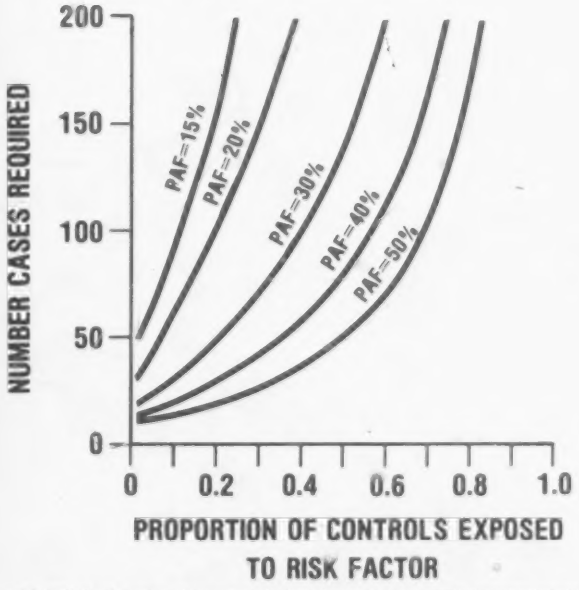

FIGURE 4—Number of cases required to detect a given population attributable fraction (PAF) as a function of the proportion of exposed controls (P_x), at control: case ratio = 2:1, power = 0.80, one-tailed alpha = 0.025.

alence of the exposure and the relative risk: for a given population attributable fraction, an increase in exposure prevalence is accompanied by a compensatory decrease in the relative risk. This decrease in the relative risk reduces power more than the increase in the prevalence raises it. The net result is a reduction in the power to detect a given population attributable fraction.

Thus only very large studies have adequate power to determine whether prevalent exposures are associated with substantial population attributable fractions. Consider a common exposure such as dietary consumption of red meat (say $P_x = 0.8$). Even a large case-control study (with 150 cases and 300 controls) would have less than 80 percent power to detect, for example, whether eating red meat accounted for 50 percent of colon cancer in the US. There are no easy solutions to this problem. Power can be somewhat increased by treating an exposure as a continuous or categorical variable if that is an option; the population attributable fraction can be determined for such exposures.[19] To maximize PAF, a continuous risk factor, such as serum cholesterol, should be dichotomized at the threshold above which disease risk appears to increase. Among the cohort of 361,662 men screened for the Multiple Risk Factor Intervention Trial, for example, dichotomizing serum cholesterol at 181 mg/dl (4.68 mmol/l), resulted in a PAF of 48 percent.[20] Alternatively, it may be necessary to compare the prevalence of

certain exposures in areas of high and low disease incidence, such as the US and Japan in the example of colon cancer,[21] to suggest potential risk factors.

By contrast, even studies with relatively few subjects often have adequate power to detect small population attributable fractions associated with a rare exposure like diabetes. Our results should be reassuring to investigators concerned about having sufficient power to detect exposures that are important from the public health point of view, but rare.

As is the case when sample size and power are based on the relative risk, these estimates assume that there are no confounders of the relationship between the exposure and the disease. Especially if the associations between exposure and confounder, and between confounder and disease, are strong, it is prudent to make separate sample size estimates at each level (stratum) of the confounder.[22,23]

Use of the relative risk as a basis for calculating sample size and power is appropriate for most observational studies. However, many such studies are also concerned with discovering exposures that are associated with a substantial population attributable fraction and have an important public health impact.[24,25] Investigators planning such studies should be aware that very large sample sizes may be required to detect meaningful population attributable fractions for common exposures, and that rare exposures associated with substantial population attributable fractions can be detected even in small studies. Calculation of sample size and power based on the population attributable fraction may be helpful to investigators interested in the potential public health impact of the associations they hope to detect.

ACKNOWLEDGMENTS

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Call for Abstracts

APHA Late-Breaker Epidemiology Exchange Session

The Epidemiology Section will again sponsor a Late-Breaker Epidemiologic Exchange at APHA's annual meeting in Chicago, Illinois. The Exchange, to be held on Wednesday, October 25, 1989, will provide a forum for presentation of epidemiologic investigations, studies, methods, etc., which have been conceived, conducted, and/or concluded so recently that abstracts could not meet the deadline for submission to other Epidemiology Sessions. Papers submitted should report on work conducted during the last 6–12 months.

Abstracts should be limited to 200 words; no special form is required. Abstracts should be submitted to Robert A. Gunn, MD, Division of Field Services, Epidemiology Program Office, Bldg. 1, Room 5127, C08, Centers for Disease Control, Atlanta, GA 30333—telephone; (404) 639-3187—and must be received by *September 15, 1989.*

Public Health Briefs

Chronic Stressors and Maternal Depression: Implications for Prevention

SUEZANNE T. ORR, PHD, SHERMAN A. JAMES, PHD, BARBARA J. BURNS, PHD, AND BRUCE THOMPSON, PHD

Abstract: We report on the use of an instrument to measure exposure to stressors among 149 women presenting with their children for pediatric care at an urban primary care center. Overall, 38.3 percent of the women had significant levels of depressive symptoms; 71.4 percent of those in the "high stress" group had an adjusted prevalence odds ratio of 5.00 [95% CI = 2.12, 11.82]. We conclude that screening in the pediatric office is feasible for identifying women at high risk of becoming depressed. (*Am J Public Health* 1989; 79:1295–1296.)

Introduction

The prevalence of significant levels of depressive symptomatology is quite high among mothers of young children.[1,2] Most such women have regular and continuing contact with a pediatrician, thus placing the pediatrician in a position to detect maternal depression. In a prior study of mothers bringing children into an urban primary care center,[3] we found the overall prevalence of significant levels of depressive symptomatology to be around 35 percent. Maternal depression has been shown to be correlated with significant physical and mental health, and school/behavioral problems among school-aged children.[4–7] Thus, high levels of depressive symptomatology have implications for the well-being of children.

Many factors are associated with increased risk of depression among women, including exposure to psychosocial stressors, lower socioeconomic status, and single parenthood. Of these factors, exposure to psychosocial stressors may be of special importance.[8–11*] Exposure to stressors, unlike demographic factors, may be amenable to interventions that could reduce depressive symptoms. Ilfeld and others have demonstrated in large community-based studies that exposure to stressors is present *prior* to the onset of depression.[8,11,12] Thus, "screening" women for exposure to stressors could lead to the development of preventive interventions.

In this paper, we present data collected from women presenting with their children in a large pediatric practice who were administered a questionnaire to assess exposure to stressors, and an instrument to measure depressive symptoms.

Methods

Data were collected during April–May 1984 at an urban pediatric primary care practice based at Sinai Hospital in Northwest Baltimore city. Mothers of children ages 5–10 were eligible for inclusion if they presented with a child for any type of pediatric care during the study period. The age restriction was imposed because one study instrument was designed for use as part of a larger research effort to identify and reduce risks for school and behavioral problems. The refusal rate was less than 5 percent.

Exposure to stressors was assessed using the Social Environment Inventory (SEI), a self-report instrument which measures exposure to acute (i.e., life events) and chronic stressors (e.g., financial, marital, employment, housing problems).[13]

Depressive symptomatology was measured using the Center for Epidemiologic Studies Depression Scale (CES-D).[14] Contingency tables were constructed to evaluate potential associations between the independent variables and CES-D category (0-15, 16 or above), and prevalence odds ratios were calculated as a measure of the strength of associations.[15] A multiple logistic regression analysis was performed in order to evaluate the associations between each independent variable and depression (as a dichotomy), controlling for the effects of all of the other independent variables. The beta coefficients and standard errors from the logistic model were used to calculate adjusted odds ratios and 95 percent confidence intervals.

Results

During the study period, 149 women were administered the study instruments. As shown in Table 1, approximately 56 percent of the women were single parents; 84 percent were Black; 20 percent had less than a high school education.

Overall, 38.3 percent of the sample scored in the depressed range (16 or greater) on the CES-D, similar to the proportion depressed in our previous report.[3] As shown in Table 1, exposure to stressors was significantly associated with risk for depressive symptomatology. The prevalence of depression among mothers scoring in the upper third of scores on the SEI was about three times that of women scoring in the lowest or middle thirds of scores on the SEI, and the adjusted prevalence odds ratio was 5.00 (95% CI = 2.12, 11.82). Terms to represent exposure-confounder interactions were not included in the model, since there appeared to be homogeneity of effects across strata of the confounders.

Discussion

These data suggest a rather high prevalence of depressive symptomatology, especially among mothers exposed to high

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*Longfellow C, Dill D, Makosky V, Zur-Szpiro S: Stressful Life Conditions and the Mental Health of Mothers. Paper presented at the Annual Meeting of the American Psychological Association, Los Angeles, California, 1981.

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TABLE 1—Prevalence of Depressive Symptoms, and Odds Ratios by Demographic and Psychosocial Factors

Variable	N (149 total)	Percent Depressed	Unadjusted Prevalence Odds Ratio	Adjusted Prevalence Odds Ratio	95% CI for Adjusted Prevalence Odds Ratio
Marital Status					
Single, separated, divorced	82	47.6	2.72	1.62	.70–3.74
Married	64	25.0	+		
+ +					
Race					
Black	125	43.2	5.32	3.32	.84–13.09
White	24	12.5	+		
Education					
<12 yrs	29	55.2	2.34	1.82	.67–4.96
≥12 yrs	119	34.5	+		
+ + +					
Exposure to Stressors					
High	42	71.4	7.4	5.00	2.12–11.82
Mod, Low	107	25.2	+		

+ Reference category.
+ + 3 missing observations for this variable.
+ + + 1 missing observation for this variable.

levels of psychosocial stressors, the strongest predictor of depressive symptoms in the multiple logistic regression model.

An obvious limitation of these data is their cross-sectional nature. Women who have high levels of depressive symptomatology may systematically inflate self-reports of exposure to psychosocial stressors, or depression may lead to problems with family and children. Alternatively, there are sound empirical reasons to believe that the occurrence of life stress precedes the onset of depressive symptomatology.[8,11,12]

A woman who scores 10 or more on the stressor measure can be considered to be at high risk of being depressed. Most of the measured stressors are of a chronic nature, rather than transient events. It may be possible to develop primary preventive interventions that reduce exposure to some stressors or, alternately, enhance the skills or resources of the women to cope with psychosocial stressors. In this connection, prior research has shown that as the number of stressors to which a woman was exposed increased, so did the risk of depression.[8] One might infer then that reducing the number of stressors to which a woman was exposed could lead to a reduction in risk of depression. Such interventions might include referrals for assistance with child care, or housing, or other such programs. Working to improve the social conditions of mothers and children may be an effective means to reduce their exposure and/or vulnerability to socioenvironmental stressors and associated higher risks of maternal depression. This, in turn, may lead to decreased risk of physical and mental illness and school problems among children.

APPENDIX

The Social Environment Inventory (SEI)

The SEI is a self-report instrument which measures exposure to acute (i.e., life events) and chronic stressors (e.g., financial, marital, employment, housing problems) (It also has items to measure access to social support, but these were not included in the present analysis). The instrument was designed to be administered in the pediatric primary care setting to women with school age children, ages 5–10. It has been tested for both reliability and validity. The reliability assessment focused on reproduceability of scores over a 30-day period. Spearman's rank order correlation coefficient over the 30-day time period for exposure to stressors was .74. The validity assessment focused on construct validity. A measure of construct validity was depression, since exposure to stressors among women is known to be associated with depression. It was found that high levels of exposure to stressors as measured by the SEI was significantly associated with an increased risk of depression (prevalence ratio = 3.7, $p < .001$). Additional analyses were conducted to determine if SEI stressor scores were associated with race or maternal education. Neither of these factors was significantly associated (using the chi square statistic for analysis) with SEI scores, suggesting that something other than those sociodemographic variables examined explained variation in scores.

The number of affirmative responses (i.e., experiencing a stressor) on the SEI are summed into a total "stress" score. In keeping with earlier analyses, scores were divided into approximately equal tertiles. For this analysis, the women with scores in the upper third (score of >10) were compared with those in the lower two-thirds.

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Immunity to Diphtheria and Tetanus in Inner-City Women of Childbearing Age

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Abstract: This study was conducted to determine immunity to diphtheria and tetanus in 232 inner-city women experiencing a recent birth. Forty-three (18.5 percent) of the women had levels of diphtheria antitoxin below the protective level (<0.01 unit/ml), whereas only 10 (4.3 percent) had insufficient levels of tetanus antitoxin. The percent of women susceptible increased with age, with 33 percent and 25 percent of women over the age of 30 years susceptible to diphtheria and tetanus, respectively. (*Am J Public Health* 1989; 79:1297–1298.)

Introduction

Tetanus and diphtheria are uncommon diseases in the United States.[1] However, both are associated with high morbidity and mortality and are preventable by immunization. For infants and children, immunization against diphtheria and tetanus is routine and mandatory for entrance into grade school. Surveys based on household interviews indicate that 95 percent of children entering grade school receive three or more doses of diphtheria-tetanus-pertussis (DTP) vaccine.[2]

Adults who have completed a primary series of tetanus and diphtheria toxoids are recommended to receive a booster every 10 years.[3] However, from 1977–82, 56 percent of diphtheria cases and 89 percent of tetanus cases occurred in adults indicating that many adults have insufficient levels of antitoxin.[3] Serosurveys in adults have indicated that tetanus and diphtheria antitoxin levels were deficient in 20 percent to 80 percent,[4] with percentages higher for women than men, especially in the middle-aged and elderly populations.[4–6]

Few recent immune status data are available on women of childbearing age from low socioeconomic areas. During a study of the immune response of preterm and fullterm infants to DTP vaccine[7] approved by the Joint Committee on Clinical Investigation of the Johns Hopkins Medical Institutions, we obtained samples from women whose infants attended one of two inner-city pediatric clinics. As a part of this study, the authors determined the immunity to diphtheria and tetanus in this group of women experiencing a recent birth.

Methods

Sera were collected between August 1984 and June 1986 from women whose infants attended one of two outpatient pediatric clinics serving mainly a Medicaid population at the Johns Hopkins Hospital.[7] The mothers of all preterm infants born during the study period and mothers of fullterm infants chosen randomly on the following clinic day were approached for participation. If the mother chosen was not present in the clinic or refused to participate, then another mother was chosen at random from the appointment list. Informed consent was obtained and a specimen of blood collected by venipuncture. Each women was interviewed and information collected on age, race, and education.

Levels of diphtheria and tetanus antitoxin were determined by the indirect hemagglutination assay adapted to microtiter technique.[8] To allow comparisons with previous studies, an antitoxin level of ≥0.01 unit/ml was considered protective.

The relation between age and percent of women susceptible was analyzed using the Chi-square test for trend.[9]

Results

One hundred fifty-seven mothers of preterm infants and 273 mothers of fullterm infants were eligible for the study; samples from 99 (63.1%) mothers of preterm infants and 133 (48.7%) mothers of fullterm infants were examined.

The study population consisted of 232 women aged 13 to 46 years (mean = 23 years). Thirty-two percent of the women were under the age of 20. Nearly all of the women were Black (95%); 117 (50%) had completed at least a high school education.

Forty-three (18.5%) of the women had levels of diphtheria antitoxin below the protective level (Table 1). There was an increasing percentage of women susceptible to diphtheria with increasing age. There were no differences in the percentage of women unprotected by level of education.

Ten (4.3%) of the women had levels of tetanus antitoxin less than 0.01 unit/ml, thus considered unprotected (Table 1). The percent of women with titers below the protective level, again, increased with age, with none of the women under the age of 20 unprotected compared to 25 percent of those over the age of 30 years. Again, there were no differences in the percent of women unprotected by level of education.

In every age group, the percent of women susceptible was greater for diphtheria than for tetanus.

Discussion

Surveys conducted in other populations of women over

TABLE 1—Number and Percent of Childbearing Women Susceptible* to Diphtheria and Tetanus by Age

Age (years)	N	Diphtheria* N (%)	Tetanus* N (%)
<16	25	2 (8.0)	0 (0.0)
17–19	49	6 (12.2)	0 (0.0)
20–22	64	12 (18.8)	1 (1.6)
23–25	36	5 (13.9)	0 (0.0)
26–30	34	10 (29.4)	3 (8.8)
>30	24	8 (33.3)	9 (25.0)
Total	232	43 (18.5)	10 (4.3)

*Susceptible = <0.01 unit/ml of antitoxin.
χ^2 test for trend for diphtheria = 7.88; p = 0.005
χ^2 test for trend for tetanus = 20.18; p < 0.0001

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the last 20 years have found a higher percent of women susceptible compared to the present study.[4,5,10] However, these studies have also reported an increase in the proportion susceptible with age. Millian, *et al*,[4] found low levels of tetanus antitoxin (<0.01 unit/ml) in 25 percent of women aged 20–29 years and 37 percent of women aged 30–39 years. Levels of diphtheria antitoxin <0.02 unit/ml were found in 15 percent of women aged 15–19 years, 19 percent of women 20–29 years, and 24 percent of women 30–39 years.[4]

The participants in this study were a group of self-selected women whose infants attended outpatient clinics that serve mainly a Medicaid population in a large city. However, they were a high-risk population for low utilization of medical care and discontinuity of care. Additional data on immunization history would have helped to further characterize this population.

In the present study, the percent of women susceptible to diphtheria was higher in all age groups compared to tetanus. One reason for this observation may be the use of tetanus toxoid alone as the practice for wound prophylaxis after trauma instead of the combined diphtheria-tetanus toxoids. Another reason is that diphtheria toxoid is less antigenic than tetanus toxoid and therefore the duration of seroprotection following immunization may be less than that observed for tetanus toxoid.[5]

The level of 0.01 unit/ml was chosen as the indicator for protection in this study to enable comparisons to other studies. However, the protective level of antitoxin for diphtheria is less well defined than that for tetanus.[11,12] Outbreaks do occur in communities in which immunization has been extensive and immunization with diphtheria toxoid may modify disease rather than prevent its occurrence.[13]

Susceptibility to diphtheria in the older age groups can be a result of lack of exposure to diphtheria in the past, lack of completion of the primary series in childhood, or lack of boosters since childhood. Since immunity to tetanus is entirely dependent on immunization, the women in the older age groups did not complete their primary series, or did not receive recent boosters. In Maryland, the requirement for immunization before entry into schools became effective in 1973 for children in nursery school through sixth grade.[14] Therefore, the women born after 1962, those under age 23 at the midpoint of the study, would have been the most likely to have been immunized in this program.

Two major areas of concern arise from these data:

- As these women age, they will contribute to the already susceptible elderly population. Second, the protection of infants against diphtheria and tetanus in their first few months of life is dependent on the mother passing adequate levels of antitoxin to the fetus during pregnancy. Given that diphtheria and tetanus antibodies readily cross the placenta during pregnancy,[15,16] regular boosters of women would help to protect their newborn infants in the first few months of life.
- All of the women in this study had been hospitalized for delivery a few months before blood drawing. The immediate post delivery to preuischarge from the hospital period could have been an opportunity to provide vaccines and toxoids recommended for adult immunization. An alternative opportunity for immunization is during visits to antenatal or pediatric clinics.

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Caffeinated Beverages and Low Birthweight: A Case-Control Study

BETTE J. CAAN, DRPH, AND MARILYN K. GOLDHABER, MPH

Abstract: A case-control study, examining the effect of first trimester maternal caffeine consumption on low birthweight, was performed in the Kaiser Permanente Medical Care Program. Heavy consumption (>3 servings/day) of coffee, cola and >300 mg/day of caffeine from all sources (coffee, tea and cola) was associated with a marginally increased risk of low birthweight. Our data support previous findings of a modest effect of caffeine consumption on fetal growth. (*Am J Public Health* 1989; 79:1299–1300.)

Introduction

Several relatively small studies performed prior to a 1981 Federal Drug Administration warning which advised women to limit their caffeine intake[1] showed a positive relationship between caffeine consumption and low birth-weight (LBW) or prematurity.[2–4] Since the 1981 warning, three major studies have been reported: a retrospective study of 12,205 prenatal patients in Massachusetts showing no relation between caffeine and LBW adjusted for cigarette smoking,[5] a retrospective survey in 9,921 Japanese women demonstrating a relation between high levels of caffeine consumption and small fetal size for gestational age,[6] and a prospective study of 3,891 antenatal patients in Connecticut showing a twofold risk of LBW associated with moderate caffeine consumption and a fourfold risk of LBW associated with heavy caffeine consumption.[7]

Methods

From a large case-control study of aerial malathion spraying and adverse pregnancy outcomes, we identified a cohort of women who had been pregnant during the period of malathion spraying in Northern California (July 1, 1981–June 30, 1982) and who had sought care at the Kaiser Permanente facility located in the sprayed county, or in either of two Kaiser Permanente facilities in demographically comparable areas outside the spray area. We retrospectively followed 9,564 pregnancies, an estimated 95 percent of the pregnancies that occurred during that time period at those facilities. Details of the methods are reported elsewhere.[8] One hundred percent of women in the larger case-control study with adverse outcomes and 20 percent of women with normal live births were selected for a post-pregnancy interview. Of those women, 85 percent completed the post-pregnancy interview, which elicited average daily consumption of caffeinated coffee, tea, and soft drinks during the first trimester.

Ninety-seven women were identified among the respondents as having infants with intrauterine growth retardation, defined as two standard deviations or more below the mean weight for gestational age.[9] An additional 34 women with live born babies weighing 2500 grams or less, but without growth retardation, were included as cases. No difference in caffeine consumption was found between these two groups, and they are combined in this paper. All 131 women are referred to as LBW. One hundred thirty-six controls were randomly selected from the remainder of normal pregnancies that resulted in live births.

Five caffeine measures were used: caffeine from all beverages, caffeine from coffee, caffeine from coffee and colas combined, number of servings of tea, and number of servings of colas. For each of these measures, women were divided into three categories based on their daily intake of caffeine: none, light-to-moderate (<300 mg or <3 servings) and heavy (>=300 mg or >=3 servings). We estimated the relative risk of LBW for the light-to-moderate and heavy consumers, each compared to the nonusers, with odds ratios (OR)[10] unadjusted and adjusted for ethnicity, alcohol intake, cigarette use, prepregnancy weight, weight gain, and parity.

Results

Several demographic and behavioral characteristics of cases and controls are summarized in Table 1. Among the cases, there were more Blacks and Asians and fewer Whites and Hispanics. Cases had a significantly lower prepregnancy weight, a significantly higher consumption of cigarettes and lower weight gain during pregnancy. In both groups, cola drinks were the most frequently consumed beverage. Heavy consumption of coffee was greater than heavy consumption of either cola or tea in both the cases and controls. Only 8.1 percent of the controls and 13.8 percent of the cases drank greater than 300 mg of caffeine daily.

No increased risk was seen with light-to-moderate consumption, with the crude or adjusted OR, but there was a suggestion of increased risk with heavy consumption of coffee, of cola, and of total caffeine (Table 2).

Discussion

We found a tripled risk of delivering a LBW infant with levels of caffeine consumption greater than 300 mg/day (OR = 2.94), but the estimate was necessarily imprecise because of the small study size (95% CI = .89, 9.65). A larger study would have permitted greater precision, but the number of controls selected was restricted by cost. Only one other study demonstrated an increased risk of LBW at such low caffeine levels, with an OR of 2.5 at daily caffeine intakes between 151–300 mg.[7] In the majority of other studies demonstrating significant associations between caffeine and birth weight, positive associations were found only at much higher levels of caffeine consumption (>500 mg/day).[3,4,6]

These findings would be distorted by recall bias if women who had pregnancies with adverse outcomes selectively overestimated their caffeine consumption. This bias appears unlikely since we found no relationship between spontaneous abortion, the primary outcome of interest in the larger study, and caffeine intake.

Since fetal growth is most rapid during the third trimester of pregnancy, one might speculate that caffeine's effect on fetal growth would be most marked during that period. We made the following assumptions about caffeine intake in the

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TABLE 1—Maternal Demographic and Behavioral Characteristics of Low Birthweight Infants and Controls

Characteristics	Cases		Controls	
	Number	Percent	Number	Percent
Total	131	100.0	136	100.0
Race				
Black	12	60.0	8	40.0
White	68	45.9	80	54.1
Hispanic	17	40.5	25	59.5
Asian	33	58.9	23	41.1
Other	1	100.0	0	100.0
Coffee (no. of 5 oz servings)				
None	68	51.9	75	55.6
>0–<3	48	36.6	52	38.5
≥3	15	11.5	8	5.9
Tea (no. of 5 oz servings)				
None	100	77.1	92	68.3
>0–<3	22	16.8	37	27.4
≥3	8	6.1	6	4.4
Cola (no. of 12 oz servings)				
None	61	46.9	72	53.3
>0–<3	62	47.7	61	45.2
≥3	7	5.4	2	1.5
Caffeine Consumption (mg/day)				
0	26	20.0	28	20.7
>0–<100	45	34.6	51	37.8
100–<200	25	19.2	29	21.5
200–<300	16	12.3	16	11.9
300–<400	7	5.4	6	4.4
400–<500	6	4.6	1	0.7
>500	5	3.8	4	3.0
	Mean	SD	Mean	SD
Prepregnancy wt (lbs)	125.32	2.54	136.43	2.49
Maternal age	26.47	.45	26.94	.44
Number of previous live births	.46	.04	.56	.04
Number of alcoholic drinks per week	.78	.18	.36	.18
Number of cigarettes per day	4.02	.67	2.01	.66
Weight gain during pregnancy (lbs)	28.13	1.24	35.14	1.19

third trimester in order to explain the biological plausibility of our hypotheses. Heavy consumers during the first trimester were likely to continue heavy consumption throughout their pregnancy, while nonconsumers either continued their same pattern of nonconsumption or increased to light-to-moderate consumption in later trimesters.

The evidence from this study, in combination with evidence from several other studies carried out over the last decade, suggests that caffeine, as found in some caffeine-containing beverages, has a small but measurable effect on fetal growth.

TABLE 2—Crude and Adjusted* Odds Ratio of Low Birthweight Associated with Exposure to Caffeinated Beverages

Caffeine Exposure	Crude Odds Ratio	Adjusted* Odds Ratio and 95% C.I.
All sources (coffee, tea + cola drinks) (mg)		
Light-moderate vs none	.96	.90 (.42–1.92)
Heavy vs none	1.76	2.94 (.89–9.65)
Coffee (no. of servings)		
Light-moderate vs none	1.02	1.02 (.53–1.95)
Heavy vs none	2.07	2.79 (.89–8.69)
Cola drinks (no. of servings)		
Light-moderate vs none	1.20	1.31 (.71–2.43)
Heavy vs none	4.13	5.87 (.72–47.83)
Tea (no. of servings)		
Light-moderate vs none	.54	.57 (.28–1.17)
Heavy vs none	1.21	1.50 (.37–6.13)
Coffee + cola drinks (mg)		
Light-moderate vs none	1.17	1.34 (.68–2.64)
Heavy vs none	2.30	3.53 (1.05–11.81)

*Adjusted for ethnicity (Black, White, Hispanic, Asian, other), alcohol intake (no. of drinks/day), cigarette use (<10 cigarettes/day, ≥10 cigarettes/day), prepregnancy weight (lbs), weight gain (lbs), parity (1, >1).

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